Data, Uncertainty and Error Analysis

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Preface

Los Angeles, CA

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This is a collection of lecture notes, problem and solutions on the topic of experimental measurements, data uncertainty and error analysis. These notes were assembled over a period of approximately four years while teaching chemistry 114 at UCLA, a physical chemistry laboratory course that includes separate lectures in addition to the laboratory sessions. The course was initially based on the excellent textbooks by J.R. Taylor, “An Introduction to Error Analysis” and Hughes & Hase, “Measurements and their Uncertainties”, but was later revised to focus on the probabilistic foundations of classical measurements. I do not consider these notes to be a substitute to the book of Taylor, which I recommend to any newcomer for its strong pedagogical value. Instead, I view these notes as providing a more in-depth coverage of the probabilistic foundation. While the reader is assumed to know calculus, no knowledge of probability theory is assumed; the required concepts are introduced as needed in these notes.

I make frequent use of MATLAB while teaching the course because it is important for young students to learn scientific computing. For those who can’t afford MATLAB, a free software alternative can be download, called GNU Octave. Many of the MATLAB examples herein should work on GNU Octave either directly or with a small amount of conversion effort. An entire chapter is dedicated to MATLAB sessions where the students are walked through several examples. At UCLA these MATLAB sessions are done during class time at the Science Learning Center. Problems and solutions
are included at the end of many chapters. The students should be aware that more advanced analysis techniques exist (such as statistical learning) and are not covered here due to the short (1-quarter) nature of this course. I would encourage the reader wanting to learn more, to read books on modern multivariate statistical techniques. Special thanks go to Alison Ly, a UCLA undergraduate student, for redrawing most of the figures in this document.

Louis Bouchard
Experimental Uncertainties

Every measurement contains some amount of uncertainty due to a variety of experimental factors (e.g. temperature fluctuations, Johnson noise, shot noise, sample motion, vibrations, environmental fields, etc.). The uncertainty is due to fluctuations in the physical quantities being measured and in the measurement apparatus.Uncertainties can be thought as representing the noise magnitude which affects our signal. The goal of the experimentalist is to reduce uncertainties to an acceptable minimum either by repeating the measurement several times or by designing a better experiment. Our goal here is to understand where uncertainties come from and how to characterize them.

1.1. Types of uncertainties

There are two main classes of uncertainties encountered by the experimentalist:

- **Systematic errors**: these errors originate because there is a bias in the system. For example, performing the same experiment on a different day could mean the ambient temperature is different, which could then lead to a drift in some currents in the system. A good piece of equipment should be designed to take into account temperature variations, for example. However, not all instrumentation is designed to compensate for environmental factors. Another example could be the measurement of a magnetic field using a magnetometer. However, the presence of large metal objects nearby could affect the magnetic field.
1. Experimental Uncertainties

- **Random errors:** these errors arise from random fluctuations in the electronics or the physical measurement under study. For example, if you are measuring the voltage across the terminals of a load, there will be random fluctuations in the voltage as function of time. These could be due to Johnson noise, which is due to thermal fluctuations of electrons in the resistance of the load and leads to random voltage fluctuations. Johnson noise is also generally present in the measuring apparatus. Vibrations could also give rise to random errors.

Note: Sometimes the distinction between systematic and random error may depend on time scales. The systematic errors can be randomly fluctuating quantities that change so slowly that they appear static on the timescale of (rapid) measurements. Random errors are those due to fluctuations that are rapid compared to the timescale of measurement. In the latter case, a series of consecutive measurements appears to fluctuate randomly.

1.2. Histograms

The quantities measured in the laboratory are random variables. A random variable is not a variable in the usual sense. It is instead characterized by a probability distribution function that encodes all relevant statistical information about the random variable.

Suppose that $x$ is a random variable. For example, let $x$ be the diameter of CDs produced in a factory. If we example $n$ CDs and measure their diameters, we collect $n$ measurements of $x$: 

\[
\{x_1, x_2, \ldots, x_n\}.
\]

These values are all different because no two CDs are perfectly identical. $x$ fluctuates; so $x$ is therefore a random variable. Suppose that we measure 10,000 CDs; we would get a list of diameter such as:

119.73  
117.10  
122.76  
119.20  
119.12  
117.94  
122.35  
121.26  
120.97  
122.66  
.
.
.

Figure 1.1. Histogram is a discrete approximation of the probability distribution of a random variable. It is obtained from experimental data.

(A list of 10,000 values in total.) The nominal diameter of a CD is 120 mm.

Now suppose that we we plot a histogram of these $x$ values as follows: bin the $x$ axis into 20 bins (i.e. from 110 to 130), with each bin of width 1. These bins could be centered on those values, for example: $[109.5, 110.5],$ $[119.5, 120.5],$ $[120.5, 121.5],$ $[121.5, 122.5],..., etc. For each bin, count the number of times $x$ falls within the interval of the bin, i.e. # of times you find a value in the above list that falls within that interval. Obviously, the longer the list (the larger $n$ is), the larger this count will be.

Plot this value (frequency of occurrence) vs $x$, as shown in Fig. 1.1. This histogram is a (discrete) approximation of the probability density function (PDF), which is a function that describes the distribution function of the random variable $x$. The finer the bins are, and the larger $n$ is, the more closely this histogram approximates a continuous function.

The histogram is also called empirical distribution. Mathematically, we partition the horizontal axis in $N$ “bins”, defined by the intervals $(r_k, r_{k+1}]$, where $\cdots < r_0 < r_1 < r_2 < \ldots$, $r_i \in \mathbb{R} \cup \{-\infty, +\infty\}$. Let $X_1, X_2, \ldots, X_n$ be the data points to be binned and $n$ the number of data points. The histogram can be formally defined as the function:

$$\hat{h}(x) = \frac{1}{n} \sum_{k=1}^{N} \frac{\mathbb{I}(r_k, r_{k+1})(x)}{(r_{k+1} - r_{k})} \sum_{i=1}^{n} \mathbb{1}\{X_i \in (r_k, r_{k+1}]\}$$

\footnote{This plot was generated in GNU Octave using the following sequence of commands: \texttt{x=120+2*randn([10000 1])}; and \texttt{figure; hist(x,linspace(110,130,21))}; The first command creates a “fake” data set that simulates the acquisition of experimental data with random error (more specifically, a random variable $x$ with mean 120 plus normally-distributed random noise with standard deviation of 2).}
where \(1_{\{X_i \in (r_k, r_{k+1}]\}}\) is an indicator function that equals 1 when \(X_i \in (r_k, r_{k+1}]\) and 0 otherwise. Similarly, \(I_{(r_k, r_{k+1})}(x)\) is an indicator function that equals 1 when \(x \in (r_k, r_{k+1}]\) and 0 otherwise. \(\sum_{i=1}^{n} 1_{\{X_i \in (r_k, r_{k+1}]\}}\) counts the number of times a result \(X_i\) falls into the bin \((r_k, r_{k+1}]\). The coefficient \(\sum_{k=1}^{N} I_{(r_k, r_{k+1})}(x)\) ensures that \(x \in (r_k, r_{k+1}]\) (\(k\) is a counter that loops over all bins, one at a time). In Problem 27, we prove that the histogram converges to the PDF in the limit of large \(n\).

### 1.3. Methods for reporting uncertainties

#### 1.3.1. Method 1.

When reporting a measured quantity \(x\) and its uncertainty \(\delta x\), we write

\[
(\text{measured value of } x) = x \pm \delta x
\]

For example, the length of a rod is measured to be \((1.0 \pm 0.1)\) m. In this course we will present common ways to estimate the value of \(\delta x\) and which value of \(x\) to report. In a certain sense, this notation means that most values measured experimentally will fall within the interval \([x - \delta x, x + \delta x]\).

In reality, however, the measured value of \(x\) is the sum of the true value of \(x\), \(x_{\text{true}}\), systematic errors (bias) plus a random error (\(\xi\)):

\[
(\text{measured value of } x) = x_{\text{true}} + \text{bias} + \xi
\]

where \(\xi\) is a random variable whose outcomes fall within the interval \([-\delta x, \delta x]\) to a large extent. If the value of the bias is known, it can always be subtracted from the measurement to obtain \(x_{\text{true}} \pm \delta x\). In some cases it is possible to design experiments to specifically measure bias.

#### 1.3.2. Method 2 (notation).

Another method for reporting uncertainties uses brackets to list the digits that are uncertain:

\(1.234(55)\) m is a more-compact way of writing \(1.234 \pm 0.055\) m.

#### 1.3.3. Report your uncertainties with 1 or 2 significant figures.

Experimental uncertainties should be rounded to 1 or 2 significant figures. For example,

\((10.000 \pm 0.123)\) m

should really be rounded to

\((10.0 \pm 0.1)\) m

or

\((10.00 \pm 0.12)\) m.
Two significant figures are normally used in precision measurements. In most other cases, we keep only one significant figure. In this course, we will stick to 1 figure because we are not doing precision measurements.

We round the uncertainty to the 1 significant figure and then report the value of \( x \) to the same digits. For example, writing

\[
(30.129 \pm 0.1) \text{ m}
\]

does not make sense because the number of significant figures to the right of the period differ for each number. We should instead report

\[
(30.1 \pm 0.1) \text{ m},
\]

where the value of \( x \) has been rounded to the same decimal place as the error.

In most calculations, it is advisable to keep many significant figures during the calculation and only round off at the end, when the final uncertainty has been determined. This is to avoid round-off errors.

When comparing two different reported values, we can determine if these values differ significantly or not by looking at the interval defined by the uncertainties. For example, if someone reports two different lengths,

\[
(10 \pm 5) \text{ m and } (18 \pm 5) \text{ m},
\]

we see that the error bars overlap: the first interval is

\[ [5, 15] \]

while the second is

\[ [13, 23] \]

Since the error bars overlap, we cannot say for sure that the two numbers differ significantly. On the other hand if the numbers reported are

\[
(10 \pm 1) \text{ m and } (18 \pm 1) \text{ m},
\]

which correspond to intervals

\[ [9, 11] \text{ and } [17, 19] \]

In this case, we can say that the two values are significantly different from each other.

Figure 1.2 illustrates the error bars for the case where resistance A is measured to be \((15 \pm 1) \Omega \) and resistance B is measured to be \((25 \pm 2) \Omega \) versus the case where resistance C is measured to be \((16 \pm 8) \Omega \) and resistance D is \((26 \pm 9) \Omega \).

Since the error bars do not overlap, values A and B are said to be significantly different from each other. Whereas measurements C and D do not differ significantly because of the substantial overlap of their error bars. Even
1. Experimental Uncertainties

Figure 1.2. Discrepancy between two measurements. We must look at the errors bars, not just the nominal value. (left) No overlap (right) Some overlap. In the second case, we cannot say for sure that the two measured values are significantly different because of the overlap.

![Figure 1.2](image)

Figure 1.3. Histograms tell us about accuracy (if the true value is known) and precision. The latter is related to the spread of the distribution.

![Figure 1.3](image)

though the $x_{best}$ values differ by 10, in one case, the values are significantly different whereas in the other case they are not.

1.4. Random Errors, Systematic Errors and Mistakes

There are three main types of errors encountered in experiments:

- **Random errors**: These errors are random fluctuations that influence the precision (or “spread”) of the measurement, as shown in Fig. 1.3. The width of the distribution provides a measure of the random error. The common method for reducing random errors is by repeating the measurement many times and taking the average. We will see later that this
1.5. Uncertainty of a single measurement

Suppose that we want to determine the uncertainty of a single measurement, so we can write \( x_{\text{best}} \pm \delta x \), where \( x_{\text{best}} \) here is the results of a single measurement. The random error can be obtained as follows:

- **From the measurement device:** If you are using a digital device, use the manufacturer’s specs or a full last digit, e.g.
  \[ (1.56 \pm 0.01) \, \Omega \]
  For analog devices, report half of a division, e.g.
  
  \[
  \text{measure 8.5 m} \quad \rightarrow \quad \text{report (8.50 \pm 0.05) m}
  \]
Here, there is no problem adding a zero to 8.5 because with the analog device, we are able to “eyeball” the extra digit by estimating the position between the two divisions. An example of analog device is shown in Fig. 1.4.

It is also possible to estimate the uncertainty yourself if the divisions are large enough. In that case, you can estimate the last digit based on the position between two divisions and estimate the uncertainty in your procedure. For example, the analog device may only have divisions every 1 cm. But you may be able to estimate up to the millimeter by eyeballing the measurement. This procedure may not necessarily give you precision to the 1 mm scale, but you may be able to estimate it within 3 mm. Suppose the measurement is 5 cm according to the divisions available on the analog device, and you estimate the last digit to be 5.3 cm. You would then report 5.4 ± 0.3 cm.

- **Using the standard deviation:** The standard deviation can be used to represent the error in a single measurement. The problem is that we need many measurements to obtain the standard deviation, i.e. we need to repeat the measurement $n$ times and compute the “sample standard deviation”

$$
\sigma_{n-1} = \sqrt{\frac{1}{n-1} \sum_{i=1}^{n} (x_i - \bar{x})^2} \quad \text{(sample standard deviation)}
$$

where $\bar{x}$ is the mean value of $x$, which may be estimated as the arithmetic sum

$$
\bar{x} = \frac{1}{n} \sum_{i=1}^{n} x_i \quad \text{(sample mean)}
$$

This quantity is called the “sample mean” because it is the calculation of the mean from the samples $x_1, \ldots, x_n$. Other accepted names for these terms include “population standard deviation” and “population mean”.\(^2\)

Note: One problem with the second method is that the measurement precision may be limited by the finite resolution of the measurement device. For example, it is possible that repeated measurements all give the same answer:

$$10.0 \text{ m}, 10.0 \text{ m}, 10.0 \text{ m}, 10.0 \text{ m}, \ldots, 10.0 \text{ m}.$$ 

If that is the case, use the first method (estimate the error bar from the smallest division of the measurement device.)

\(^2\)It may seem ridiculous to collect $n$ measurements just to compute the error bar of a single measurement; true, but it is nonetheless the correct way to obtain the error in a single measurement, when repeated measurements are possible. If you can afford to perform $n$ measurements, report the mean value instead; in which case, the error in the mean is called the standard error, which is defined below.
1.6. Standard error

If you report for $x_{\text{best}}$ a mean value calculated from repeated measurements, the error in the measurement of $x_{\text{best}}$ is the standard error, not the standard deviation. The standard error is defined as the ratio of the sample standard deviation to $\sqrt{n}$

$$\alpha = \frac{\sigma_{n-1}}{\sqrt{n}}$$

If we collect $n$ data points, each with uncertainty $\sigma_{n-1}$ ("sample standard deviation" of a single measurement), and calculate the sample mean $\bar{x}$, the uncertainty in the sample mean is $\alpha$.

We report

$$\bar{x} \pm \alpha = \bar{x} \pm \frac{\sigma_{n-1}}{\sqrt{n}}$$

The standard error is also called the "standard deviation of the mean".

Note: the "standard error" is defined as above, in terms of the standard deviation. However, the concept of error bars being reduced by $\sqrt{n}$ still applies to any error bar when the measurements are averaged. For example, if your error bar in a single measurement of $x$ is $\delta x$, as obtained (say) from the manufacturer specs of your measurement instrument, then this error will be reduced to $\delta x/\sqrt{n}$ when reporting the average $\bar{x}$ from $n$ measurements as calculated by the arithmetic mean. Technically, this is not the standard error because we use $\delta x$ instead of $\sigma$, but it is a highly related concept.

1.7. Reporting experimental results for a mean value

When doing experiments, we typically repeat a measurement many times to reduce its random error. The number reported is the sample mean. (Even if, as an experimentalist, you don’t do these repeats yourself, it is likely that your measurement apparatus does this averaging for you.) Here is how this procedure is handled (either manually by you or automatically by your experimental apparatus):

- Collect experimental data as $n$ samples $x_1, x_2, \ldots, x_n$.
- Calculate the sample mean $\bar{x}$ as the arithmetic sum $\bar{x} = \frac{1}{n} \sum_{i=1}^{n} x_i$ while keeping all the significant figures.
- Calculate the standard error (error in the mean $\bar{x}$), $\alpha = \frac{\sigma_{n-1}}{\sqrt{n}}$ where

$$\sigma_{n-1} = \sqrt{\frac{1}{n-1} \sum_{i=1}^{n} (x_i - \bar{x})^2}.$$
1. Experimental Uncertainties

while keeping all the significant figures.

- Decide how many significant figures to retain for the error. If many data points are used, keep 2 significant figures. Otherwise keep only 1. In this course, we will stick to retaining only 1 significant figure for the error bar.
- Round the mean to the appropriate decimal place.

**Reporting errors:**

If the experimental measurement can be repeated many times (e.g., $x_1, x_2, \ldots, x_n$), the most common way to report the experimental value of $x$ is:

$$\text{(measured } x) = \bar{x} \pm \frac{\sigma_{n-1}}{\sqrt{n}}$$

where $\bar{x} = \frac{1}{n} \sum_{i=1}^{n} x_i$ and $\sigma_{n-1} = \sqrt{\frac{1}{n-1} \sum_{i=1}^{n} (x_i - \bar{x})^2}$. This assumes that $\sigma_{n-1}$ is a meaningful (non-zero) value, i.e. it is constructed from non-identical measurements (such as 10.0, 10.0, 10.0, 10.0, etc).

If it is not possible to repeat the experimental measurement (due to time constraint or other reason), then report the single measured data point, $x_1$, plus uncertainty as stated by the manufacturer or using a pre-calculated value for the standard deviation.\(^a\)

\(^a\)Although calculating the standard deviation does require repeated measurements, this task can be done at a more convenient time.

1.8. Precision vs. Accuracy

Precision and accuracy are different concepts.

1.8.1. Accuracy. Suppose you know the “true value” of a physical quantity. For example, the speed of light in vacuum is a universal physical constant that is defined to be equal to

$$c = 299,792,548 \text{ m/s}.$$  

This figure is exact since the length of the meter is defined from this constant and the international standard for time. The “true value” can be obtained from other means. The accuracy is a measure of how close is the mean value of your measurement from the true value.

1.8.2. Precision. Loosely speaking, precision is defined as the “spread” or scatter of values around the mean. A precise measurement corresponds to a small spread, whereas an imprecise measurement corresponds to a large spread.
1.8. Precision vs. Accuracy

In Fig. 1.5, four distinct cases are illustrated: accurate and precise measurement, accurate and imprecise measurement, inaccurate but precise measurement and inaccurate and imprecise measurement. The graphs shown are obtained by plotting histograms of the frequency of a given experimental result versus the value measured.

If the “true value” corresponds to the dotted line, an accurate measurement of $x$ means that the average value of the measurement, which can be thought of as the “center of mass” of the histogram, lies close to the dotted line. This is true, regardless of the sharpness or spread around this mean value. An inaccurate measurement is one where the measured values cluster far away from the dotted line.

1.8.3. Fitting data to a model. As scientists, we often need to validate theory with experiments. Physical quantities are related to one another via some physical law (a mathematical formula). In which case, there is an equation (model) available to fit your data to. By fitting data to your
model, you can see if the model suitably describes the physical situation. (And there are situations, of course, where the model may not be known and analysis of the data requires you to identify a suitable model.)

Experimental data comes with error bars. Plotting error bars on a graph enables us to decide if experimental data is consistent with a given model. Let us look at the example of Hooke’s law which describes the extension of a spring from its equilibrium position when attaching a mass to its end.

Example: Hooke’s law. Hooke’s law says that the force $F$ on the spring is linearly proportional to the displacement $x$ from equilibrium ($x = 0$):

$$F = kx$$

If we attach a mass $m$ to the spring, the force acting on the spring is the gravitational force on the mass $m$:

$$F = mg,$$

where $g$ is the gravitational acceleration constant, approximately $g \sim 9.8 \text{ m/s}^2$. The situation is illustrated in the figure below. Writing the Hooke’s law in the form

$$x = \frac{F}{k} = \frac{mg}{k} = \left( \frac{g}{k} \right)m$$

immediately suggests a possible experiment for measuring the value of $k$: we attach different masses $m$ and measure the corresponding displacement $x$. The slope of this graph yields $g/k$, from which we can obtain $k$ (Fig. 1.6). Figure 1.6 also shows an example where linear fit does not describe the data well and a quadratic component must be added to the model.

### 1.8.4. Fractional Uncertainties.

It is sometimes convenient to express values

$$(\text{measured } x) = x_{\text{best}} \pm \delta x$$

in terms of fractional uncertainty

$$(\text{fractional uncertainty}) = \frac{\delta x}{|x_{\text{best}}|}$$

If we multiply by 100 we get the percentage uncertainty:

$$(\text{percentage uncertainty}) = \frac{\delta x}{|x_{\text{best}}|} \times 100$$

If we multiply by 1,000,000 we get “parts per million”:

$$(\text{parts per million (ppm)}) = \frac{\delta x}{|x_{\text{best}}|} \times 10^6.$$  

and similarly for “parts per billion” (ppb). The main advantage of fractional uncertainty is that it is a dimensionless quantity (it has no units).
1.9. Rules for Rounding and Significant digits

1.9.1. Significant Digits. Here are accepted rules for reporting significant digits

- All non-zero digits are significant:
  
  \[2.998 \times 10^8 \text{ m/s}\] has 4 significant figures

- All zeroes between non-zero digits are significant
  
  \[6.022 \, 141 \, 79 \times 10^{23} \text{ mol}^{-1}\] has 9 significant figures

- Zeroes to the left of the first non-zero digits are not significant
  
  \[0.51 \text{ MeV}\] has 2 significant figures

- Zeroes at the end of a number to the right of the decimal point are significant
  
  \[1.60 \times 10^{-19} \text{ C}\] has 3 significant figures
• If a number ends in zeroes without a decimal point, the zeroes might be significant:

\[ 270 \, \Omega \] could have 2 or 3 significant figures

To avoid confusion, report instead:

\[ 2.70 \times 10^2 \, \Omega \] or \[ 2.7 \times 10^2 \, \Omega \]

1.9.2. Rounding rules. How to round numbers:

• If the next digit is 4 or lower, leave unchanged:

\[ 6.62 \times 10^{-34} \] becomes \[ 6.6 \times 10^{-34} \]

if 2 significant figures are appropriate.

• If the next digit is 6 or higher, increase by 1:

\[ 5.67 \times 10^{-8} \] becomes \[ 5.7 \times 10^{-8} \]

• If the digit after last one to be retained is 5
  – Leave last digit unchanged if even:

\[ 9.45 \] becomes \[ 9.4 \] (2 sig figs)

since the (circled) digit 4 is even
  – Increase last digit by 1 if it is odd:

\[ 9.55 \] becomes \[ 9.6 \] (2 sig figs)

since the (circled) digit 5 is even.

Rounding the result of addition and subtraction: We round-off the result to the same number of decimal places as the number with the least number of decimal places:

\[ 1.23 + 45.6 = 46.8 \]

(We have dropped the '3' in 1.23 because 45.6 does not contain such a level of precision; hence the '3' becomes meaningless when adding.)

1.9.2.1. Rounding the result of multiplication and division. We keep the same number of significant figures as the component with the least number of significant figures:

\[ \underline{1.2} \times \underline{345.6} = 414.72 = \underline{4.1} \times 10^2 \]

For more complicated operations than addition/subtraction or multiplication/division of two numbers, we will be learning methods of error propagation later in the course.
1.10. Problems

Problem 1. Partial differentiation: (a) Find the total differential of the function \( f(x, y) = y \exp(x + y) \). (The total differential of \( f(x, y) \) is defined as \( df = \frac{\partial f}{\partial x} dx + \frac{\partial f}{\partial y} dy \).)

(b) Find the first and second partial derivatives of the function \( f(x, y) = 2x^3y^2 + y^3 \), i.e. calculate \( \frac{\partial f}{\partial x}, \frac{\partial f}{\partial y}, \frac{\partial^2 f}{\partial x^2}, \frac{\partial^2 f}{\partial y^2}, \frac{\partial^2 f}{\partial x \partial y}, \frac{\partial^2 f}{\partial y \partial x} \).

(c) Suppose we have a function \( f(x, y) \) where \( y = y(x) \). The total derivative of \( f \) with respect to \( x \) is

\[
\frac{df}{dx} = \frac{\partial f}{\partial x} + \left( \frac{\partial f}{\partial y} \right) \frac{dy}{dx}.
\]

Find the total derivative of \( f(x, y) = x^2 + 3xy \) with respect to \( x \), given that \( y = \sin^{-1} x \).

Solution. (a) \[
\frac{\partial f}{\partial x} = y \exp(x + y), \quad \frac{\partial f}{\partial y} = \exp(x + y) + y \exp(x + y).
\]

\[
df = [y \exp(x + y)]dx + [(1 + y) \exp(x + y)]dy.
\]

(b) \[
\frac{\partial f}{\partial x} = 6x^2y^2, \quad \frac{\partial f}{\partial y} = 4x^3y + 3y^2, \quad \frac{\partial^2 f}{\partial x^2} = 12xy^2, \quad \frac{\partial^2 f}{\partial y^2} = 4x^3 + 6y,
\]

\[
\frac{\partial^2 f}{\partial x \partial y} = 12x^2y = \frac{\partial^2 f}{\partial y \partial x}
\]

(c) \[
\frac{\partial f}{\partial x} = 2x + 3y, \quad \frac{\partial f}{\partial y} = 3x, \quad \frac{dy}{dx} = \frac{1}{(1 - x^2)^{1/2}}
\]

and so

\[
\frac{df}{dx} = 2x + 3y + 3x \frac{1}{(1 - x^2)^{1/2}} = 2x + 3 \sin^{-1} x + \frac{3x}{(1 - x^2)^{1/2}}.
\]

Problem 2. Find the gradient of the following functions:

(a) \( r = \sqrt{x^2 + y^2 + z^2} \)

(b) \( f(x, y, z) = x^2 + y^3 + z^4 \)

(c) \( f(x, y, z) = x^2y^3z^4 \)

(d) \( f(x, y, z) = e^x \sin(y) \log(z) \)

Solution. (a) \( \nabla r = \mathbf{r} \). (Since \( \partial_x r = (1/2)(1/r)2x = x/r, \) etc.)

(b) \( \partial_x f = 2x, \ \partial_y f = 3y, \ \partial_z f = 4z \). So that \( \nabla f = (2x, 3y, 4z) \).
The Taylor expansion of a function of two variables is:

\[ y \text{ and } x \]

\[ \nabla f = (2xy^3z^4, 3x^2y^2z^4, 4x^2y^3z^3). \]

So that

\[ \nabla f = (2xy^3z^4, 3x^2y^2z^4, 4x^2y^3z^3). \]

(d) \( \partial_x f = e^x \sin(y) \log(z), \) \( \partial_y f = e^x \cos(y) \log(z), \) \( \partial_z f = e^x \sin(y)z^{-1}. \)

So that \( \nabla f = e^x(\sin(y) \log(z), \cos(y) \log(z), \sin(y)z^{-1}). \)

Problem 3. For the three questions below let \( r - r' \) be the separation vector from a fixed point \( (x', y', z') \) to the point \( (x, y, z) \) and \( |r - r'| \) be its length. If \( \nabla \) indicates derivative with respect to the unprimed variables \( r \), show that:

(e) \( \nabla|r - r'|^2 = 2(r - r') \)

(f) \( \nabla(1/|r - r'|) = -(r - r')/|r - r'|^3 \)

(g) What is the general formula for \( \nabla(|r - r'|^n)? \)

(h) In problem (f) above you have computed \( \nabla(1/|r - r'|) = -(r - r')/|r - r'|^3. \)

Now calculate the quantity \( \nabla'(1/|r - r'|), \) where \( \nabla' \) denotes the derivative with respect to the primed variables \( r' \) (instead of the unprimed variables \( r \)).

Solution. For (e), the \( x \) component of the gradient is \( \partial_x((x - x')^2 + (y - y')^2 + (z - z')^2) = 2x, \) etc. so that \( \nabla|r - r'|^2 = 2(r - r'). \) For (f) the \( x \) component is \( \partial_x 1/\sqrt{(x - x')^2 + (y - y')^2 + (z - z')^2} = (1/2)((x - x')^2 + (y - y')^2 + (z - z')^2)^{-3/2}(2x) \) so that \( \nabla(1/|r - r'|) = -(r - r')/|r - r'|^3. \) For (g) we have \( \nabla|r - r'|^n = n(r-r')/|r-r'|^{n-1}. \) For (h) the only difference is a negative sign.

Problem 4. Show that the Taylor expansion, up to quadratic terms in \( x - 2 \) and \( y - 3, \) of \( f(x, y) = y \exp(xy) \) about the point \( x = 2, y = 3, \) is:

\[ f(x, y) = e^6\left\{ 3 + 9(x - 2) + 7(y - 3) + (2!)^{-1}\left[ 27(x - 2)^2 + 48(x - 2)(y - 3) + 16(y - 3)^2 \right] \right\} \]

(Show all your calculations)

Solution. The Taylor expansion of a function of two variables is:

\[ f(x+h, y+k) = f(x, y) + (h, k) \cdot \nabla f(x, y) + \frac{1}{2} [h \ k] \left[ \begin{array}{cc} \partial^2 f/\partial^2 x^2 & \partial^2 f/\partial x \partial y \\ \partial^2 f/\partial y \partial x & \partial^2 f/\partial^2 y^2 \end{array} \right] \left[ \begin{array}{c} h \\ k \end{array} \right] + O(|(h, k)|^3) \]

To obtain the stated result, plug in \( x = 2 \) and \( y = 3 \) in the Taylor expansion formula and take \( h = x - 2 \) and \( k = y - 3 \) as the step size, making use of the following derivatives (evaluated at the point \( x = 2 \) and \( y = 3 \)):

\[ \frac{\partial f}{\partial x} = y^2 \exp(xy), \quad \frac{\partial f}{\partial y} = \exp(xy) + xy \exp(xy), \quad \frac{\partial^2 f}{\partial x^2} = y^3 \exp(xy), \]

\[ \frac{\partial^2 f}{\partial y \partial x} = \exp(xy), \quad \frac{\partial^2 f}{\partial^2 y^2} = \exp(xy). \]
\[ \frac{\partial^2 f}{\partial y^2} = 2x \exp(xy) + x^2 y \exp(xy), \quad \frac{\partial^2 f}{\partial x \partial y} = 2y \exp(xy) + xy^2 \exp(xy). \]

Problem 5. Shorthand notation: \( f_x \equiv \frac{\partial f}{\partial x}, \quad f_{xy} = \frac{\partial^2 f}{\partial x \partial y}, \quad f_{xx} = \frac{\partial^2 f}{\partial x^2}, \) etc. The multivariable Taylor expansion can be used to study the behavior of functions near extrema. All stationary points have \( f_x = f_y = 0 \) and these points may be classified as:

1. minima if both \( f_{xx} \) and \( f_{yy} \) are positive and \( f_{xy}^2 < f_{xx} f_{yy} \),
2. maxima if both \( f_{xx} \) and \( f_{yy} \) are negative and \( f_{xy}^2 < f_{xx} f_{yy} \),
3. saddle point if \( f_{xx} \) and \( f_{yy} \) have opposite signs or \( f_{xy}^2 > f_{xx} f_{yy} \).

Prove that the function \( f(x,y) = x^4 + y^4 \) has a minimum at the origin, but that \( g(x,y) = x^4 + y^3 \) has a saddle point there.

Solution. We have \( f_x \equiv \partial_x f = 4x^3 \) and \( f_y \equiv \partial_y f = 4y^3 \), both of which are zero at the origin. Thus the origin is a stationary point. To check that this is a maximum, minimum or not, we need the second derivatives: \( f_{xx} = 12x^2 \), \( f_{yy} = 12y^2 \), \( f_{xy} = 0 \). Both \( f_{xx} \) and \( f_{yy} \) are zero at the origin, so the test is inconclusive. However, both are positive in a small neighborhood of the origin and \( f_{xy}^2 < f_{xx} f_{yy} \). Thus, we have a minimum.

Problem 6. Show that the function \( f(x,y) = x^3 \exp(-x^2 - y^2) \) has a maximum at the point \((\sqrt{3}/2, 0)\), a minimum at \((-\sqrt{3}/2, 0)\). What about the nature of the point at the origin?

Solution. Setting the first two partial derivatives to zero to locate the stationary points, we find

\[ \frac{\partial f}{\partial x} = (3x^2 - 2x^4) \exp(-x^2 - y^2) = 0, \quad \frac{\partial f}{\partial y} = -2xy^3 \exp(-x^2 - y^2) = 0. \]

For the second equation to be satisfied we require \( x = 0 \) or \( y = 0 \) and for the first one to be satisfied we require \( x = 0 \) or \( x = \pm \sqrt{3}/2 \). Hence the stationary points are at \((0,0)\), \((\sqrt{3}/2, 0)\) and \((-\sqrt{3}/2, 0)\). We now find the second partial derivatives:

\[ f_{xx} = (4x^5 - 14x^3 + 6x) \exp(-x^2 - y^2) \]
\[ f_{yy} = x^3(4y^2 - 2) \exp(-x^2 - x^2) \]
\[ f_{xy} = 2x^2 y(2x^2 - 3) \exp(-x^2 - y^2) \]

We then substitute the pairs of values of \( x \) and \( y \) for each stationary point and find that at \((0,0)\)

\[ f_{xx} = 0, \quad f_{yy} = 0, \quad f_{xy} = 0 \]
1. Experimental Uncertainties

and at \((\pm \sqrt{3/2}, 0)\)

\[ f_{xx} = \mp \sqrt{3/2} \exp(-3/2), \quad f_{yy} = \mp \sqrt{3/2} \exp(-3/2), \quad f_{xy} = 0. \]

Hence applying the above three criteria, we find that \((0, 0)\) is an undetermined stationary point, \((\sqrt{3/2}, 0)\) is a maximum and \((-\sqrt{3/2}, 0)\) is a minimum. The function is shown in Figure 1.7.

\[ (x^3 \exp(-x^2-y^2)) \]

\[ x \quad 0 \quad 1 \quad 2 \]
\[ y \quad -3 \quad -2 \quad -1 \quad 0 \quad 1 \quad 2 \quad 3 \]

Figure 1.7. Plot of the function \(f\).

**Problem 7.** Matrix inverse. (a) Derive the formula for matrix inverse of a 2x2 matrix:

\[
A^{-1} = \begin{bmatrix} a & b \\ c & d \end{bmatrix}^{-1} = \frac{1}{\det(A)} \begin{bmatrix} d & -b \\ -c & a \end{bmatrix} = \frac{1}{ad-bc} \begin{bmatrix} d & -b \\ -c & a \end{bmatrix}.
\]

(b) Check that it satisfies the *definition* of the matrix inverse, namely check that \(A^{-1}A = AA^{-1} = I\) where \(I\) is the 2 \times 2 unit matrix.

(c) Prove that taking the inverse of \(n \times n\) matrices \(A, B\) reverses the order of matrices:

\[(AB)^{-1} = B^{-1}A^{-1}\]

(d) Prove that \((k\text{ is a non-zero scalar})\)

\[(kA)^{-1} = k^{-1}A^{-1}\]

where \(A\) is an invertible matrix.

(e) Prove that taking the inverse of an invertible matrix twice recovers the original matrix:

\[(A^{-1})^{-1} = A\]

**Problem 8.** Manipulation of matrices and matrix operations. (a) For two arbitrary matrices \(A\) and \(B\), write down the *matrix products* \(AB\) and \(BA\). What are the conditions on \(A\) and \(B\) for the matrix product to *exist* (and be well-defined)?

(b) In general, does \(AB\) equal \(BA\)?
(c) Check that for matrices $A, B$ and scalar $c$ the following property holds: 
\[(A^T)^T = A\]
where $A^T$ denotes matrix transpose.

(d) Prove that the transpose operation preserves the matrix addition: 
\[(A + B)^T = A^T + B^T\]

(e) Prove that taking the transpose of a product of matrices reverses the orders of the matrices 
\[(AB)^T = B^T A^T\]

(f) Prove that a scalar is left invariant by the transpose operation: 
\[(cA)^T = cA^T\]

(g) Prove that the dot product of two column vectors $a$ and $b$ can be computed as 
\[a \cdot b = a^T b\]
where $a^T$ denotes the transpose of $a$ (i.e. a row vector). Verify this by writing out explicitly the matrix product.

(h) Prove that the transpose of an invertible matrix is also invertible and its inverse is the transpose of the inverse of the original matrix: 
\[\left( A^T \right)^{-1} = (A^{-1})^T\]

Problem 9. Compute the following dimensionless quantity and find the correct error bars (pay attention to the order of operations indicated by the brackets): 
\[
\left[ (0.0045 \pm 0.0005) \text{ plus } (0.3 \pm 0.9) \right] \text{ divided by } (1.5 \pm 0.1) = ?
\]

Solution. From the first operation (addition) we have 
\[\alpha_Z = \sqrt{(\alpha_A)^2 + (\alpha_B)^2} = \sqrt{(0.0005)^2 + (0.9)^2} = 0.9\]
and 
\[Z_{best} = 0.0045 + 0.3 = 0.3045\]
which we round according to the error $= 0.3045 \pm 0.9 = 0.3 \pm 0.9$. Then we perform the second operation (division). There, we have 
\[Z'_{best} = \frac{0.3}{1.5} = 0.2\]
and the error bar 
\[\frac{\alpha_{Z'}}{Z'} = \sqrt{\left( \frac{\alpha_A}{A} \right)^2 + \left( \frac{\alpha_B}{B} \right)^2}\]
evaluates to
\[ \alpha Z' = 0.2 \sqrt{\left(\frac{0.9}{0.3}\right)^2 + \left(\frac{0.1}{1.5}\right)^2} = 0.6. \]

And the final result is 0.2 ± 0.6.

**Problem 10.** Suppose that a random experiment consists of measuring the length \( L \) of an object. (a) Explain the possible sources of randomness in this measurement.

(b) The “state space” of this random experiment consists of all possible values \( L \) can take (without regards to the way \( L \) is measured, i.e. assume infinite precision). Is this state space continuous or discrete?

(c) Random variables depend on how they are defined relative to an experiment. Define \( X \) as the random variable which consists of the value \( L \) measured with infinite precision. Is \( X \) a continuous or discrete random variable?

(d) Let \( Y \) be the random variable which corresponds to the length \( L \) rounded to the precision of the measuring instrument. Is \( Y \) a discrete or continuous random variable?

(e) Let \( Z \) be the random variable which equals 1 if \( L \) is greater than 10 mm and 0 if \( L \) is less than or equal to 10 mm. (Such random variables are called “indicator functions”, because they act as logical bits which are “on” when certain conditions are met and “off” otherwise.) Is \( Z \) a discrete or continuous random variable?

**Solution.** (a) Any experiment performed in the lab suffers from random errors (fluctuations). Their origin could be any of: rapid thermal fluctuations, mechanical vibrations, operator (the act of measurement can introduce errors), fluctuations in the electrical power supply, etc.

(b) Continuous. (Since the values of \( L \) are not denumerable.)

(c) Continuous. (\( X \) can take an indenumerable number of values.)

(d) Discrete. (The number of values \( Y \) can take can be infinite, but still denumerable.

(e) Discrete, since \( Z \) can only take 2 possible values: 0 and 1.

**Problem 11.** Let \( f \) be a differentiable real-valued function on \( \mathbb{R}^3 \) and \( v_p \) a tangent vector of \( \mathbb{R}^3 \) at a point \( p \in \mathbb{R}^3 \). Then
\[ v_p[f] = \frac{d}{dt} (f(p + tv))|_{t=0} \]
to the directional derivative is the (directional) derivative of \( f \) with respect to \( v_p \).

For example, suppose that \( f = x_1x_2x_3 \), \( p = (1, -4, 2) \) and \( v = (1, 1, 0) \), where \( x_1, x_2, x_3 \) are the coordinate functions of \( \mathbb{R}^3 \).
(a) Find \( p + tv \) and show that \( f(p + tv) = 2(1 + t)(-4 + t) \). Show that \( v_p[f] = -6 \) by direct computation of the limit.

(b) If \( v_p = (v_1, v_2, v_3) \) is a tangent vector of \( \mathbb{R}^3 \) at a point \( p \), then

\[
v_p[f] = \sum_{i=1}^{3} v_i \frac{\partial f}{\partial x_i}(p).
\]

Prove this statement from the above definition (limit).

(c) Use the definition of directional derivative in (b) to show that \( v_p[f] = -6 \).

(d) Let \( f \) and \( g \) be differentiable functions on \( \mathbb{R}^3 \), \( v_p \) and \( w_p \) tangent vectors at a point \( p \), and \( a \) and \( b \) real numbers. Prove the following 3 properties:

(i) \( (av_p + bw_p)[f] = av_p[f] + bw_p[f] \),

(ii) \( v_p[af + bg] = av_p[f] + bv_p[g] \),

(iii) \( v_p[fg] = v_p[f] \cdot g(p) + f(p) \cdot v_p[g] \).

**Solution.**

(a)

\[
\begin{align*}
p + tv &= (1 + t, -4 + t, 2), \\
f(p + tv) &= 2(1 + t)(-4 + t) \\
v_p[f] &= \frac{d}{dt}2(1 + t)(-4 + t)|_{t=0} = 2(2t - 3)|_{t=0} = -6.
\end{align*}
\]

(b) Let \( p = (p_1, p_2, p_3) \). Then,

\[
f(p + tv) = f(p_1 + tv_1, p_2 + tv_2, p_3 + tv_3).
\]

Since \((d/dt)(p_i + tv_i) = v_i\), by putting \( x_i = p_i + tv_i \) and using the chain rule we obtain the desired result.

(c)

\[
\begin{align*}
\frac{\partial f}{\partial x_1}(p) &= x_2x_3(p) = -8, \\
\frac{\partial f}{\partial x_2}(p) &= x_1x_3(p) = 2, \\
\frac{\partial f}{\partial x_3}(p) &= x_1x_2(p) = -4.
\end{align*}
\]

we obtain

\[
v_p[f] = 1(-8) + 1(2) + 0(-4) = -6.
\]

**Problem 12.** Report the following numerical distance correctly, with error bars: \( X + Y \), where \( X = 110.125 \pm 0.003 \) m and \( Y = 85.6 \pm 0.2 \) m.

**Solution.** Whether we use the \( \sigma_Z = \sigma_X + \sigma_Y \) method or the quadrature method, the error in \( X + Y \) is dominated by the error in \( Y \) (i.e. \( \sigma_Y \gg \sigma_X \)). Thus the error bar is 0.2 m. Next, we round \( 110.125+85.6 = 195.725 \) to the tenths digit, giving \( (195.7 \pm 0.2) \) m.
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Problem 13. How many significant figures are there in this expression: 3000000000 liters.

Solution. Anywhere between 1 and 10.

Problem 14. Express the following result in proper rounded form, with suitable error bars: mass=19.1234 g with uncertainty 0.6789 g.

Solution. First we need to round the error bar to 1 or 2 sig figs. Let’s do 1 sig fig: 0.7 g. Then we round the mass to this figure: 19.1 g. The result is: 19.1(7) g.

Problem 15. Your experiment yielded the following measurement: (4.1234 ± 0.4321) Joules. Report this number with proper error bars and appropriate significant figures.

Solution. Either (4.1 ± 0.4) J or (4.12 ± 0.43) J.

Problem 16. How many significant figures are there in each of the following expressions?

(i) 0.00082 s  
(ii) 0.14800 psi  
(iii) 6.24×10^6 l  
(iv) -754.090×10^{-27} J  
(v) 50 cm  
(vi) 50 m

Solution. (i) 2  
(ii) 5  
(iii) 3  
(iv) 6  
(v) 1 or 2  
(vi) 1 or 2

Problem 17. Express the following result in proper rounded form, with suitably truncated error bars: mass=8.4857 g with uncertainty 0.2554 g.

Solution. First we need to truncate the error bar to 1 or 2 sig figs. Let’s do 1 sig fig: 0.3 g. Then we round the mass to this figure: 8.5 g. The result is: 8.5(3) g.

Problem 18. You measure the length of an object with a ruler (or measuring tape) whose smallest division is 1 mm. You measure the length 5 times with results in mm of 123.4, 123.5, 124.6, 123.7, 123.8 mm (the last
digit you have estimated by eyeballing). What is the average length and the uncertainty in length?

**Solution.** Because this is an “analog” device (ruler’s smallest division is 1 mm), we should take 1/2 or the smallest division as our error. Namely, 0.5 mm. Thus, the result is 123.8(5).

**Problem 19.** Compute the following dimensionless quantity and find the correct error bars (pay attention to the order of operations indicated by the brackets):

\[
[(0.0045 \pm 0.0005) \text{ plus } (0.3 \pm 0.9)] \text{ divided by } (1.5 \pm 0.1) =?
\]

**Solution.** From the first operation (addition) we have

\[
\alpha_Z = \sqrt{(\alpha_A)^2 + (\alpha_B)^2} = \sqrt{(0.0005)^2 + (0.9)^2} = 0.9
\]

and

\[Z_{\text{best}} = 0.0045 + 0.3 = 0.3045\]

which we round according to the error = 0.3045 ± 0.9 = 0.3 ± 0.9. Then we perform the second operation (division). There, we have

\[Z'_{\text{best}} = \frac{0.3}{1.5} = 0.2\]

and the error bar

\[
\frac{\alpha_{Z'}}{Z'} = \sqrt{\left(\frac{\alpha_A}{A}\right)^2 + \left(\frac{\alpha_B}{B}\right)^2}
\]

evaluates to

\[
\alpha_{Z'} = 0.2 \sqrt{\left(\frac{0.9}{0.3}\right)^2 + \left(\frac{0.1}{1.5}\right)^2} = 0.6.
\]

And the final result is 0.2 ± 0.6.
Experimental measurements in the laboratory are random variables (rv). Every time you measure a physical quantity you get a different number because of random fluctuations (random errors). The random fluctuations are called random errors. Thus, a statistical description of experimental measurements is needed.

Here, we introduce tools to study random variables. Random variables can be continuous or discrete, depending on whether they take continuous or discrete values, respectively. An example of a continuous random variable is the length of an object. Length is a random variable which can take positive real values in a continuous interval. An example of a discrete random variable is the number of counts within a time interval. Counts can only take discrete values \((1, 2, 3, \ldots)\), in this case, the positive integers.

2.1. Continuous Random Variables

A random variable \(X\) is not a simple variable. \(X\) is instead better described by a function. This function encodes all of its statistical properties.\(^1\) We associate with \(X\) a probability density function (PDF), \(p(x)\).

As a matter of convention, we shall use capital letters, e.g. \(X\), to denote random variables and lowercase letters, e.g. \(x\), for the value taken by the random variable \(X\).

\(^1\)Think of an experiment performed on Monday. The value measured on Tuesday may be slightly different than the one obtained on Monday because of random errors. Same story for measurements performed on subsequent days — these values will also be different due to fluctuations.
2.2. Probability Density Function

The probability density function (PDF) of a random variable $X$, denoted $p(x)$, is everywhere non-negative: $p(x) \geq 0$ and is normalized to 1:

$$\int_{-\infty}^{\infty} p(x) dx = 1.$$

We note that the PDF refers to a particular random variable (say, $X$). This is sometimes emphasized by adding a subscript, e.g., $p_X(x)$ instead of $p(x)$. When working with a single random variable we do not need the subscript because it should be clear that there is only one possible random variable that $p(x)$ refers to. However, when the problem involves more than one random variable, we should use a subscript to avoid confusion between the different PDFs.

2.3. Cumulative Distribution Function

We define the cumulative distribution function (CDF) as the integral of the PDF:

$$P(X \leq x) \equiv \int_{-\infty}^{x} p(x') dx'.$$

The CDF is the probability that $X$ takes a value less than or equal to $x$. The quantity $\{X \leq x\}$ is an example of a random event; the function $P(\cdot)$ associates a number between 0 and 1 to this random event. We note that if $p(x)$ is continuous, then there is no distinction between $P(X \leq x)$ and $P(X < x)$. When discontinuities are present, we should be careful about the equality.

From this definition, we can solve for $p$ in terms of $P$: \(^2\)

$$p(x) = \frac{dP(X \leq x)}{dx} \bigg|_{x}.$$

We note that

$$P(a \leq X \leq b) = \int_{a}^{b} p(x) dx = \left( \int_{-\infty}^{b} - \int_{-\infty}^{a} \right) p(x) dx = P(X \leq b) - P(X \leq a).$$

2.3.1. Interpretation of PDF. The PDF, $p(x)$, has the following interpretation. The quantity $p(x)dx$ is the probability that the random variable $X$ lies in the interval $[x, x + dx]$:

$$p(x)dx = P(x \leq X \leq x + dx),$$

\(^2\)To differentiate the integral with respect to $x$, apply the Leibniz formula for differentiation of integrals. In the expression $\int_{-\infty}^{x} p(x') dx'$, the only dependence on $x$ comes from the upper limit of the integral. Thus, $\frac{d}{dx} \int_{-\infty}^{x} p(x') dx' = p(x)$. 

2. Probability
where \( dx \) is an infinitesimally small quantity and \( \mathbb{P}(\cdot) \) denotes the probability of the event \( \cdot \) occurring. The quantity \( p(x)dx \) by itself is rarely used, except under the integral sign. Instead, one integrates this expression to find the probability that \( X \) will take some value in a finite interval \([a, b] \):

\[
\mathbb{P}(a \leq X \leq b) = \int_a^b p(x)dx.
\]

The last step follows from

\[
\mathbb{P}(x \leq X \leq x + dx) = \mathbb{P}(X \leq x + dx) - \mathbb{P}(X \leq x) = d\mathbb{P}(X \leq x),
\]

integrating \( \mathbb{P}(x \leq X \leq x + dx) = d\mathbb{P}(X \leq x) \) from \( a \) to \( b \) yields \( \int_a^b d\mathbb{P}(X \leq x) = \mathbb{P}(X \leq b) - \mathbb{P}(X \leq a) = \mathbb{P}(a \leq X \leq b) \) whereas integrating \( p(x)dx \) yields \( \int_a^b p(x)dx \). Since the two are equal, we have that

\[
\mathbb{P}(a \leq X \leq b) = \int_a^b p(x)dx.
\]

2.3.2. Limit value of CDF. We note that the CDF approaches 1 in the limit of large \( x \). This follows from the normalization condition on the PDF.

2.4. Mean Value of Continuous Random Variable

Let \( X \) be a continuous random variable. The mathematical expectation of \( X \), denoted \( \mathbb{E}(X) \), is defined as:

\[
\mathbb{E}(X) \equiv \int_{-\infty}^{\infty} x p(x)dx
\]

where the integral is over all values taken by the random variable \( X \) (here, over the entire real line). If the random variable takes values in the interval \([0, 1]\) then the limits of the integral range from 0 to 1.

That is, to obtain the mean value of \( X \), we replace the random variable \( X \) by a regular variable \( x \) that represents its value, then multiply by \( p(x) \) and integrate over all \( x \).

Other names for \( \mathbb{E}(X) \) include “mean value” (of \( X \)), or “expectation value” or “average value”. Other symbols you may encounter in the literature include \( \overline{X} \), \( \langle X \rangle \) or \( m(X) \).

We note that this expression differs from the sample mean (see previous lecture): \( \overline{X} = \frac{1}{n}\sum_{i=1}^{n} x_i \). The sample mean is only an estimate of the mean.\(^4\) We will derive the formula for the sample mean later in the course.

\(^3\)In calculus, recall that \( df(x) = f(x + dx) - f(x) \).

\(^4\)More specifically, the sample mean is the best estimate of the mean in the sense of least squares.
2.5. Indicator Functions

An indicator function, \(1_{\{x<y\}}(x)\), is a function that takes the value 1 when \(x < y\) and 0 otherwise. Indicator functions can also be applied to random events. Let \(X\) be a random variable and \(A\) is a random event. The indicator function for the random event \(X \in A\) is defined as:

\[
1_A(X) = \begin{cases} 
1 & \text{if } X \in A \\
0 & \text{otherwise}
\end{cases}
\]

where \(X \in A\) refers to the value of the rv \(X\) taken after a random experiment. Another notation for \(1_A(X)\) you may encounter is \(1_{X \in A}\). You may also encounter \(I_{X \in A}\) instead of \(1_{X \in A}\). Taking the mathematical expectation of \(1_{X \in A}\) and applying the definition of probability,

\[
\mathbb{E}1_{X \in A} = \int_{-\infty}^{\infty} 1_{x \in A}(x) p(x) dx = \int_{A} p(x) dx = \mathbb{P}(A).
\]

where \(A\) is a random event and the integral \(\int_{A}\) means integral over all points \(x\) that meet the condition \(x \in A\) (for example, \(X\) could be a coordinate, and \(A = (-\infty, y]\) indicates an event where the coordinate is less than \(y\)). Indicator functions are useful when dealing with experimental measurements. See Problems 27, 22 and 28 for example uses of the indicator function.

2.6. Variance

The variance of \(X\), denoted \(\text{var}(X)\) or \(\sigma^2_X\), is defined as the sum of square differences between \(X\) and its mean, \(\mu_X \equiv \mathbb{E}X\), weighted by the PDF:

\[
\sigma^2_X \text{ or } \text{var}(X) \equiv \int_{-\infty}^{\infty} p(x)(x - \mu_X)^2 dx
\]

The square can be expanded to give \(\int_{-\infty}^{\infty} p(x)(x^2 + \mu^2_X - 2x\mu_X) dx\) and thus

\[
\sigma^2_X = \mathbb{E}(X^2) - (\mu_X)^2.
\]

The square root of the variance is called the standard deviation and is commonly denoted \(\sigma\).

2.7. Example PDFs

2.7.1. Cauchy Distribution. Let \(X\) be a random variable with the Cauchy (or Lorentzian) distribution. Its PDF is defined as:

\[
p_X(x) = \frac{1}{\pi} \frac{1}{1 + x^2}.
\]
The CDF is:

\[ P(X < x) = \int_{-\infty}^{x} \frac{1}{\pi} \frac{1}{(1 + x^2)} \, dx. \]

We know from calculus that the derivative of \( \tan^{-1}(x) \) is \( 1/(1 + x^2) \). Therefore, the last expression can be integrated:

\[ P(X < x) = \frac{1}{\pi} \tan^{-1}(x) + \frac{1}{2}. \]

### 2.7.2. Rayleigh Distribution.

The Rayleigh distribution has a PDF:

\[ p(x) = \frac{x}{\sigma^2} e^{-x^2/(2\sigma^2)}, \]

where \( x \geq 0 \) and \( \sigma \) is a parameter of the distribution. The CDF can be shown to be (\( X \sim \text{Rayleigh} \)):

\[ P(X < x) = 1 - e^{-x^2/(2\sigma^2)}, \]

where \( x \geq 0 \). The reader can check that the mean of \( X \) is \( \sigma \sqrt{\pi}/2 \) and its variance is \( \sigma^2 4\pi - 2 \).

### 2.7.3. Gaussian (Normal) Distribution.

The normal distribution \( N(\mu, \sigma^2) \) with parameters \( \mu \) and \( \sigma^2 \) is defined by the density:

\[ p(x) = \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{(x-\mu)^2}{2\sigma^2}}, \]

where \( x \in (-\infty, \infty) \). The prefactor \( \frac{1}{\sqrt{2\pi\sigma^2}} \) is such that \( p(x) \) adds up to 1:

\[ \int_{-\infty}^{\infty} p(x) \, dx = 1. \]

This can be verified using the well-known result for a Gaussian integral \( \int_{-\infty}^{\infty} e^{-ax^2} \, dx = \sqrt{\pi/a} \), where \( a > 0 \). As an exercise, you should check that this PDF is normalized to 1, the mathematical expectation of \( X \sim N(\mu, \sigma^2) \) is \( \mu \) and its variance is \( \sigma^2 \), i.e. \( \mathbb{E}X = \mu \) and \( \mathbb{E}(X - \mu)^2 = \sigma^2 \).

This probability density is plotted below. It is centered on \( \mu = \mathbb{E}(X) = \bar{X} \) and the width is proportional\(^5\) to \( \sigma \).

If a random variable \( X \) follows a Gaussian distribution (Fig. 2.1) with mean \( \mu \) and variance \( \sigma^2 \) we write \( X \sim N(\mu, \sigma^2) \). For a Gaussian distribution, the CDF is called the error function. See Figure 2.2.

\(^5\)In fact, the full width at half maximum of the Gaussian is \( 2\sqrt{2\log 2} \sigma \approx 2.355 \sigma \). You can check this by finding the values of \( x \) for which \( \frac{1}{2} \cdot \frac{1}{\sqrt{2\pi\sigma^2}} = \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{x^2}{2\sigma^2}} \) or \( \frac{1}{2} = e^{-\frac{2\sigma^2}{2\sigma^2}} \), since the maximum of the function is \( \frac{1}{\sqrt{2\pi\sigma^2}} \) (set \( x = 0 \)). Taking logs of both sides gives \( x = \pm \sqrt{2\sigma^2 \log 2} \).
\( \text{Erf}(x)_{\mu,\sigma} = \frac{1}{\sqrt{2\pi\sigma^2}} \int_{-\infty}^{x} e^{-\frac{(x-\mu)^2}{2\sigma^2}} \, dx. \)

The error function is an integral which cannot be evaluated analytically. Instead it must be solved numerically. Values of the error function can be obtained from tables, calculators or computer programs. The error function for standard normal rv (mean 0, variance 1) is often tabulated in books. In MATLAB the command \texttt{normcdf(x,mu,sigma)} will return values for Erf\((x)_{\mu,\sigma}\). See Section 2.8 for a discussion of the error function.

2.7.4. Log-Normal Distribution. In Section 2.7.3, we have introduced the error function as the CDF of the Gaussian PDF (Eq. 2.1):

\[ \text{Erf}(x)_{\mu,\sigma} = \frac{1}{\sqrt{2\pi\sigma^2}} \int_{-\infty}^{x} e^{-\frac{(x-\mu)^2}{2\sigma^2}} \, dx. \]

If \( X \sim N(\mu, \sigma^2) \) and \( Y = e^X \), then

\[ \mathbb{P}(Y < y) = \mathbb{P}(e^X < y) = \mathbb{P}(X < \log y), \]
2.8. Tabulated Values of Error Function

which leads to the CDF:

\[ P(Y < y) = \int_{-\infty}^{\log y} \frac{1}{\sqrt{2\pi\sigma^2}} e^{-(x-\mu)^2/2\sigma^2} dx = \text{Erf}(\log y)_{\mu,\sigma}. \]

This is called the log normal distribution \((Y \sim \text{log-normal})\). You can check, using the Leibniz formula for differentiation, that the PDF of the log-normal distribution is:

\[ p_Y(y) = \frac{1}{\sqrt{2\pi\sigma^2}} e^{-(\log(y)-\mu)^2/2\sigma^2} \cdot \frac{1}{y}. \]

2.8. Tabulated Values of Error Function

Let \(X \sim N(\mu, \sigma^2)\). Then,

\[ \phi_{\mu,\sigma}(x) \equiv \text{Erf}(x)_{\mu,\sigma} = P(X < x) = \frac{1}{\sqrt{2\pi\sigma^2}} \int_{-\infty}^{x} e^{-(\xi-\mu)^2/(2\sigma^2)} d\xi \]

let \( \zeta=(\xi-\mu)/\sigma, d\xi=d\zeta/\sigma \)

\[ = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{(x-\mu)/\sigma} e^{-\zeta^2/2} d\zeta = \Phi \left( \frac{x-\mu}{\sigma} \right), \]

where \( \Phi(\cdot) \) denotes the normalized error function:

\[ \Phi(z) = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{2} e^{-\zeta^2/2} d\zeta. \]

The latter is the CDF of the standard normal distribution, \(N(0,1)\). \(z\) is known as the \(z\)-score:

\[ z = \frac{x-\mu}{\sigma}. \]

As an example, Eq. (2.2) can be expressed in this notation as:

\[ \text{Erf}(\log y)_{\mu,\sigma} = \Phi \left( \frac{\log y - \mu}{\sigma} \right). \]

You should beware that there exist other conventions for the error function. For example, MATLAB and EXCEL softwares define the error function as:

\[ \text{erf}(x) = \frac{2}{\sqrt{\pi}} \int_{0}^{x} e^{-t^2} dt. \]

This is related to the normal CDF according to:

\[ \phi_{\mu,\sigma}(x) = \text{Erf}(x)_{\mu,\sigma} = \Phi \left( \frac{x-\mu}{\sigma} \right) = \frac{1}{2} \left( 1 + \text{erf} \left( \frac{x-\mu}{\sigma\sqrt{2}} \right) \right). \]

This expression can be used to calculate \(\text{Erf}(x)_{\mu,\sigma}\) using data from tables, where \(\mu\) is the mean of the population and \(\sigma\) is the standard deviation of the
population. First, we form the “z-score” (Eq. 2.3). Then we use tabulated values of the error function for this particular value of $z$.

For example, suppose that a manufacturer produces electrical resistors whose nominal value is $(100 \pm 2) \Omega$, where $2 \Omega$ is the standard deviation (both could be estimated, for example, using sample mean and sample variance). Assuming that the distribution of the resistance $X$ is Gaussian (i.e. $X \sim N(100, 2^2)$), what is the probability that choosing a resistor at random will yield a resistance of $95 \Omega$ or less? We want to show that

$$P(X \leq 95 \Omega) = \text{Erf}(95)_{100,2} \approx 0.0062.$$  

Method 1 uses MATLAB:

```matlab
>> normcdf(95,100,2)
ans =
 0.0062
```

Method 2 uses tabulated values of $\Phi(z)$: The $z$-score is:

$$z = \frac{x - \mu}{\sigma} = \frac{95 - 100}{2} = -2.5,$$

which is negative. Unfortunately, tables of error function do not list negative $z$ values. However, notice that negative $z$ values can be obtained from positive ones:

$$\Phi(-z) = 1 - \Phi(z).$$

Here, for positive $z = 2.5$ the value $\Phi(2.5)$ is 0.993790. Taking $1 - \Phi(2.5)$ gives 0.00621, the result we sought. Most books on statistics will have such a table. Tables can also be generated in MATLAB by typing:

```matlab
normcdf(linspace(0,3,50)',0,1)
```

The results $(x, \Phi(x))$ are shown in Table 2.1.

### 2.9. Confidence Limits and Error Bars

Recall the Gaussian probability density which has a bell shape centered on $\mathbb{E}X = \mu_X$ and full width at half-maximum $\approx 2.355\sigma$ (Fig. 2.3). The area under the curve bounded by the interval $x \in [\mu_X - \sigma, \mu_X + \sigma]$ is given by:

$$\frac{1}{\sqrt{2\pi\sigma^2}} \int_{\mu_X - \sigma}^{\mu_X + \sigma} e^{-\frac{(x-\mu_X)^2}{2\sigma^2}} \, dx = \text{Erf}(\mu_X + \sigma)_{\mu_X,\sigma} - \text{Erf}(\mu_X - \sigma)_{\mu_X,\sigma} \approx 0.683$$

where

$$\text{Erf}(x)_{\mu,\sigma} = \frac{1}{\sqrt{2\pi\sigma^2}} \int_{-\infty}^{x} e^{-\frac{(x-\mu)^2}{2\sigma^2}} \, dx.$$
2.9. Confidence Limits and Error Bars

Table 2.1. Numerical values of the error function \( \Phi(x) = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{x} e^{-x^2/2} \, dx \).

<table>
<thead>
<tr>
<th>( x )</th>
<th>( \Phi(x) )</th>
<th>( x )</th>
<th>( \Phi(x) )</th>
<th>( x )</th>
<th>( \Phi(x) )</th>
<th>( x )</th>
<th>( \Phi(x) )</th>
<th>( x )</th>
<th>( \Phi(x) )</th>
<th>( x )</th>
<th>( \Phi(x) )</th>
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<td>0.4856</td>
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</tr>
<tr>
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<td>0.4657</td>
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<td>0.3302</td>
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<td>0.3476</td>
</tr>
</tbody>
</table>

Figure 2.3. Confidence limits. The Gaussian distribution has a full width at half-maximum of approximately 2.355\( \sigma \).

About 2/3 of the total area under the curve is within \( \pm \sigma \) of the mean. Recall that:

\[
\text{(value of } x) = x_{\text{best}} \pm \delta x
\]

where we often take \( \delta x = \sigma \). This choice for \( \delta x \) corresponds to the “68% confidence level”. This means that we are confident, at the 68% level, that were we to take another measurement, the value would lie within one standard deviation of the mean. There are other possible conventions for choosing \( \delta x \). Common choices for \( \delta x \) are:

\[
\begin{align*}
\pm \sigma & \rightarrow \ 68\% \text{ level} \\
\pm 2\sigma & \rightarrow \ 95\% \text{ level} \\
\pm 3\sigma & \rightarrow \ 99.7\% \text{ level}
\end{align*}
\]

2.9.1. Example: From CDF to PDF. It is important to understand how this works, and not let the notation confuse you. Suppose that we have
a CDF:

\[(2.5) \quad \mathbb{P}(Y < a) = \int_{-\infty}^{a} \frac{1}{\pi} \frac{dy}{1 + y^2}\]

The PDF of the distribution is obtained by differentiating with respect to \(a\) (and renaming of \(a\) as \(y\)):

\[p_Y(y) = \frac{1}{\pi} \frac{1}{1 + y^2},\]

where we have used \(\frac{dp(X \leq x)}{dx} = p_X(x)\). Except that here we have \(a\) instead of \(x\). The differentiation is always with respect to the upper bound the statement \(\{X \leq a\}\) or \(\{X \leq x\}\). \(X\) is the random variable. Another way to look at it is to write \(F(x) = \mathbb{P}(X \leq x)\) and

\[\frac{dF(x)}{dx} = p_X(x) \quad \text{or} \quad \frac{dF(a)}{da} = p_X(a).\]

Inspection of the Leibniz formula for differentiation (see next section) shows that the differentiation step is completely trivial and amounts to simply identifying the integrand. This is consistent with the definition of CDF:

\[(2.6) \quad \mathbb{P}(Y < a) = \int_{-\infty}^{a} p_Y(y) dy.\]

2.10. Calculating Probabilities: Single Variable

We have seen in the previous lecture that probabilities of the type \(\{a \leq X \leq b\}\) are calculated by integrating the PDF from \(a\) to \(b\):

\[\mathbb{P}(a \leq X \leq b) = \int_{a}^{b} p(x) dx.\]

More generally, we deal with random events, which are statements of the type \(\{X \in A\}\) where \(A\) is a set of points. The quantity \(\mathbb{P}(X \in A)\) is a number between 0 and 1, which gives the probability for the random variable \(X\) to take the values in the set \(A\):

\[\mathbb{P}(X \in A) = \int_{\{x| x \in A\}} p(x) dx,\]

where \(\{x| x \in A\}\) is the set of points \(x\) that belong to the set \(A\). The integral is a Riemann summation over the set of points \(\{x| x \in A\}\) on the real line. This notation is useful because we can transform the statement \(\{x \in A\}\) into any equivalent statement, including one that involves a change of variables. For example, the two following statements are equivalent:

\[\{X < a\} = \{\log(X) < \log(a)\}.\]
This is useful if another random variable \( Y \) is defined as \( Y = \log(X) \). In that case, evaluating the probability of \( \{ Y < b \} \) gives the same numerical result as evaluating the probability of \( \{ X < a \} \).

### 2.10.1. Average of \( f(X) \)

The average (or mean, or expectation value) of a function \( f \) of a random variable \( X \) is defined as:

\[
\mathbb{E}(f(X)) \equiv \int_{-\infty}^{\infty} f(x)p(x)dx.
\]

### 2.11. Statistical Moments (Mean, Variance, Skewness, Kurtosis)

Take \( f(x) = x^n \) in the above formula. This gives the \( n \)-th moment of \( X \):

\[
\mathbb{E}(X^n) \equiv \int_{-\infty}^{\infty} p(x)x^n dx.
\]

The case \( n = 1 \) (first moment) is called the mathematical expectation or mean value of \( X \):

\[
\mathbb{E}(X) \equiv \int_{-\infty}^{\infty} xp(x)dx.
\]

We shall often write as shorthand:

\[
\mu_X \equiv \mathbb{E}(X).
\]

As we have seen in the previous lecture, the variance of \( X \) is defined as:

\[
\sigma^2 \equiv \int_{-\infty}^{\infty} p(x)(x - \mu_X)^2 dx,
\]

which also equals \( \sigma^2 = \mathbb{E}(X^2) - \mu^2_X \). Thus, the variance is the second moment of \( X \) minus the square of the average of \( X \). Variance is also known as the second central moment of \( X \).

The \( n \)-th central moment of \( X \) is defined as:

\[
\mathbb{E}(X - \mu_X)^n \equiv \int_{-\infty}^{\infty} p(x)(x - \mu_X)^n dx.
\]

Why are moments important? Moments describe the probability distribution. There is a theorem of mathematics that says if we know the moments of all orders, we can reconstruct the entire distribution function. You already know how to obtain the sample mean and variance. The mean is just the center of mass of the distribution whereas the variance is related to its width (about the mean).
Also of interest are the skewness (3rd moment)

\[
Skew[X] = \frac{\mathbb{E}(X - \mu_X)^3}{\left(\mathbb{E}(X - \mu_X)^2\right)^{3/2}} = \frac{\mathbb{E}(X - \mu_X)^3}{\sigma^3},
\]

and the kurtosis (4th moment):

\[
Kurt[X] = \frac{\mathbb{E}(X - \mu_X)^4}{\left(\mathbb{E}(X - \mu_X)^2\right)^2} = \frac{\mathbb{E}(X - \mu_X)^4}{\sigma^4}.
\]

The skewness measures the asymmetry of the distribution about its mean whereas the kurtosis is often used to assess by how much a distribution deviates from the bell-shape. For example, if a distribution looks like a bell shape but has much longer tails, the kurtosis will reflect this.

### 2.12. Median and Percentile

The median of a random variable \(X\) is the value of \(x\) such that

\[
\mathbb{P}(X \geq x_{50}) = \mathbb{P}(X \leq x_{50}) = \int_{-\infty}^{x_{50}} p(x)dx = \frac{1}{2}.
\]

The median is a special case of percentile. The 10-th percentile of \(X\) is the value of \(x_{10}\) such that:

\[
\mathbb{P}(X \leq x_{10}) = \int_{-\infty}^{x_{10}} p(x)dx = 0.10.
\]

The \(n\)-th percentile of \(X\) is the value \(x_n\) such that:

\[
\mathbb{P}(X \leq x_n) = \int_{-\infty}^{x_n} p(x)dx = \frac{n}{100}.
\]

#### 2.12.1. \(\langle f(X) \rangle\) is NOT the Same as \(f(\langle X \rangle)\).

**Example 1:** Suppose that the kinetic energy, \(K(v) = \frac{1}{2}mv^2\), of an object of mass \(m\) is to be calculated using experimentally measured values of the velocity \(v\). Thus, \(v\) is a random variable \(V\). Since \(v\) is a random variable, \(K(v)\) is also a random variable. You determine from experiments that the velocities, \(v\), are Gaussian-distributed around 100 m/s, with a standard deviation of 1 m/s, i.e. \(p(v) = \frac{1}{\sqrt{2\pi}}e^{-(v-100)^2/2}\). What is the average kinetic energy, \(\mathbb{E}(K(V))\)?

You expect that \(\mathbb{E}(K(V))\) should be close to \(K(100) = \frac{1}{2}m(100)^2\). However, the exact value of \(\mathbb{E}(K(V))\) will depend on the distribution \(p(v)\). We need to calculate:

\[
\mathbb{E}(K(V)) = \int_{-\infty}^{\infty} \frac{1}{2}mv^2 \frac{1}{\sqrt{2\pi}}e^{-(v-100)^2/2}dv
\]

\[
= \frac{m}{2\sqrt{2\pi}} \int_{-\infty}^{\infty} v^2 e^{-(v-100)^2/2}dv \approx \frac{m}{2} \cdot 1.0001 \cdot 10^4.
\]

So fairly close to \(K(100)\) but slightly higher.
2.12.1.2. Example 2: Suppose that \( f(\theta) = \cos(\theta) \) and \( p(\theta) = 1 \) for \( \theta \in \left[ -\frac{\pi}{2}, \frac{\pi}{2} \right] \) and \( p = 0 \) elsewhere (uniform distribution). Denote the random variable as \( \Theta \). Using \( p(\theta) \) you can easily check that \( \Theta = \int_{-1/2}^{1/2} \theta \, d\theta = 0 \). The average of \( \cos(\Theta) \) is:

\[
\cos(\Theta) = \int_{-1/2}^{1/2} \cos(\theta) \, d\theta \approx 0.9589.
\]

Note: it is not equal to 1 even though the average of \( \Theta \) is 0.

2.12.1.3. Example 3: In physics and chemistry, the notations \( \langle r \rangle \) and \( r \) are used interchangeably to denote the mathematical expectation. Consider the dipole-dipole interaction between two electric dipoles. The energy of interaction depends on \( 1/r^3 \), where \( r \) is the distance separating the two point dipoles. It is easy to show that in general, \( \langle 1/r^3 \rangle \neq \frac{1}{\langle r \rangle^3} \). \( r \) is a random variable due to molecular diffusion. It has an average value \( \bar{r} \) and deviation \( \delta r \):

\[
r = \bar{r} + \delta r,
\]

where \( \bar{r} \) is deterministic and \( \delta r \) is random with zero mean. The average of \( 1/r^3 \) is Taylor-expanded about the mean \( \bar{r} \):

\[
\langle \frac{1}{r^3} \rangle = \frac{1}{(\bar{r} + \delta r)^3} = \langle \frac{1}{\bar{r}^3} \rangle - \frac{3}{\bar{r}^4} \langle \delta r \rangle + \frac{12}{\bar{r}^5} \langle (\delta r)^2 \rangle + O(|\delta r|^3)
\]

The first term is \( 1/\bar{r}^3 \) since \( \bar{r} \) is deterministic. In the second term, \( \frac{3}{\bar{r}^4} \) can come out of the angle bracket because it is a deterministic quantity. Similarly for \( \frac{12}{\bar{r}^5} \) in the third term. Thus,

\[
\langle \frac{1}{r^3} \rangle = \frac{1}{\langle r \rangle^3} - \frac{3}{\langle r \rangle^4} \langle \delta r \rangle + \frac{12}{\langle r \rangle^5} \langle (\delta r)^2 \rangle + O(|\delta r|^3)
\]

and you can see that \( \langle \frac{1}{r^3} \rangle \) is in general different from \( \frac{1}{\langle r \rangle^3} \). Since \( \langle \delta r \rangle = 0 \) we have:

\[
\langle \frac{1}{r^3} \rangle = 1 + \frac{12}{\langle r \rangle^5} \langle (\delta r)^2 \rangle + O(|\delta r|^3).
\]

We sometimes see in the literature \( \frac{1}{\langle r \rangle^3} \) in lieu of \( \langle \frac{1}{r^3} \rangle \). This is technically incorrect. However, for small values of \( \langle |\delta r| \rangle / \langle r \rangle \), it is a good approximation.

2.13. Two (Continuous) Random Variables

If we are to compute the average of an expression that is a function of more than one random variable, we need to use the joint probability density \( p_{XY}(x, y) \), which is everywhere non-negative \( (p_{XY}(x, y) \geq 0) \) and integrates

\[\text{extra terms (nonzero)}\]

\[\text{extra terms (nonzero)}\]

\[\text{extra terms (nonzero)}\]

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\[\text{extra terms (nonzero)}\]

\[\text{extra terms (nonzero)}\]

\[\text{extra terms (nonzero)}\]
to 1:  
\[ \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} p_{XY}(x, y) dx dy = 1. \]

The joint PDF is obtained from the joint CDF analogously to the single-variable case:

\[ p_{XY}(x, y) = \frac{\partial^2}{\partial x \partial y} \mathbb{P}(X < x, Y < y). \]  

The average of a function \( g(X, Y) \) would be:  
\[ \mathbb{E}(g(X, Y)) = \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} p_{XY}(x, y) g(x, y) dx dy. \]

Given a joint probability density function, \( p_{XY}(x, y) \), the marginal density function for \( X \) is obtained by integrating over \( y \):

\[ p_X(x) = \int_{-\infty}^{\infty} p_{XY}(x, y) dy. \]

Similarly, the marginal density for \( Y \) is obtained by integrating over all \( x \):

\[ p_Y(y) = \int_{-\infty}^{\infty} p_{XY}(x, y) dx. \]

Note: you can easily check that both marginals \( p_X(x) \) and \( p_Y(y) \) are bona fide densities, i.e. nonnegative and normalized to 1.

### 2.14. Statistical Independence

The marginal density is a useful concept if you are asked to check whether or not two random variables are statistically independent. Two random variables \( X \) and \( Y \) are statistically independent if the joint probability density is written as the product of densities of each variable:

\[ p_{XY}(x, y) = p_X(x) p_Y(y), \]

where \( p_X(x) \) and \( p_Y(y) \) are the marginal densities of \( X \) and \( Y \), respectively. They can be computed from \( p_{XY}(x, y) \) by integrating.

There are at least two consequences of statistical independence that we can immediately point out. First, one concerns expectation values. Consider the average of a function \( g(X, Y) \) of two rv \( X \) and \( Y \):

\[ \mathbb{E}(g(X, Y)) = \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} p_{XY}(x, y) g(x, y) dx dy = \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} p_X(x) p_Y(y) g(x, y) dx dy. \]
If $g$ factors as a product of a function of $X$ times a function of $Y$, for example $g(X,Y) = XY$ then the expectation value of $XY$ is equal to the product of expectation values of $X$ and that of $Y$:

$$E(XY) = \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} p_X(x)p_Y(y)xy \, dx \, dy = \int_{-\infty}^{\infty} p_X(x)xdx \cdot \int_{-\infty}^{\infty} p_Y(y)ydy$$

Thus, the expectation value of a product of random variables factorizes as a product of expectation values for each random variable:

$$E(XY) = E(X) \cdot E(Y).$$

The second consequence pertains to the calculation of probabilities in general, which also factors as a product. For example, the joint probability:

$$P(X \in A, Y \in B) = \int_{\{(x,y)|x \in A, y \in B\}} p_{XY}(x,y) \, dx \, dy$$

$$= \int_{\{(x,y)|x \in A, y \in B\}} p_X(x)p_Y(y) \, dx \, dy$$

$$= \int_{\{x|x \in A\}} p_X(x) \, dx \cdot \int_{\{y|y \in B\}} p_Y(y) \, dy$$

$$= P(X \in A) \cdot P(Y \in B).$$

As a special case, take the intervals $A = (-\infty, x]$ and $B = (-\infty, y]$ and we get the result that the CDFs also factorize:

$$P(X \leq x, Y \leq y) = P(X \leq x) \cdot P(Y \leq y).$$

### 2.15. Calculating Probabilities: Two Variables

Probabilities of an event $A$ are calculated by integrating the PDF over the relevant set of points which make the event $A$ true. That is, for a single random variable $X$:

$$P(X \in A) = \int_{\{x|x \in A\}} p_X(x) \, dx,$$

where $\{x|x \in A\}$ denotes the set of all points $x$ such that $x \in A$. For example, if $A = [a, b]$ (interval), we have:

$$P(a \leq X \leq b) = \int_{a}^{b} p_X(x) \, dx.$$

If $A$ is the union of two disjoint (non-overlapping) intervals $[a, b]$ and $[c, d]$, i.e. $A = [a, b] \cup [c, d]$, then the probability of $X$ taking a value in $A$ is the
sum of two integrals:
\[ \mathbb{P}(X \in A) = \int_a^b p_X(x) \, dx + \int_c^d p_X(x) \, dx. \]

For two or more random variables we integrate the joint probability density over all such points \((x, y)\) that meet the sought criterion:
\[ \mathbb{P}(X \in A, Y \in B) = \int \int_{\{(x, y) | x \in A, y \in B\}} p_{XY}(x, y) \, dx \, dy, \]
where \(\{(x, y) | x \in A, y \in B\}\) denotes the set of all points \((x, y)\) such that \(x \in A\) and \(y \in B\).

In general, for a given set of points over which the probability needs to be calculated, we must translate what this means in terms of the upper and lower limits of integration. Let us look at some specific examples. Let \((X, Y)\) be a 2-dimensional (bivariate) rv with joint density \(p_{XY}(x, y)\). The probability that the vector \((X, Y)\) will lie in the first quadrant of the 2D plane is:
\[ \mathbb{P}(X > 0, Y > 0) = \int \int_{\{(x, y) | x > 0, y > 0\}} p_{XY}(x, y) \, dx \, dy \]
\[ = \int_0^\infty \int_0^\infty p_{XY}(x, y) \, dx \, dy. \]

Suppose again that we have a random experiment involving two rv \(X\) and \(Y\). The probability that the outcome will lie in one of the first two quadrants:
\[ \mathbb{P}(X > 0) = \mathbb{P}(X > 0, Y \in (-\infty, \infty)) = \int \int_{\{(x, y) | x > 0\}} p_{XY}(x, y) \, dx \, dy \]
\[ = \int_0^\infty \left( \int_0^\infty p_{XY}(x, y) \, dx \right) \, dy. \]

### 2.15.1. Sum of \(X^2\) and \(Y^2\).

The probability that the vector \((X, Y)\) lies outside the unit circle \(\{(x, y) | x^2 + y^2 = 1\}\) is:
\[ \mathbb{P}(X^2 + Y^2 > 1) = \int \int_{\{(x, y) | x^2 + y^2 > 1\}} p_{XY}(x, y) \, dx \, dy. \]

This can be calculated explicitly by splitting the integration domain \((\mathbb{R}^2 - \{\text{unit disc}\})\) into 4 regions (Fig. 2.4).
2.15. Calculating Probabilities: Two Variables

Figure 2.4. Partition of the region $x^2 + y^2 > 1$ into 4 regions.

\[
P(X^2+Y^2 > 1) = \int_{-\infty}^{\infty} \left( \int_{-\infty}^{-1} p_{XY}(x,y) \, dx \right) \, dy + \int_{-\infty}^{\infty} \left( \int_{1}^{\infty} p_{XY}(x,y) \, dx \right) \, dy + \int_{-1}^{1} \left\{ \int_{-\infty}^{-\sqrt{1-x^2}} + \int_{\sqrt{1-x^2}}^{\infty} \right\} p_{XY}(x,y) \, dy \, dx.
\]

Another way to calculate this would be to convert $p_{XY}(x,y)$ to polar coordinates $p_{R,\theta}(r,\theta)$ and integrate from $r = 1$ to $\infty$ while letting $\theta$ range from 0 to $2\pi$.

2.15.2. Sum of $X$ and $Y$. Here is another application of the Leibniz formula. Suppose we have two random variables $X, Y$ with joint density $p_{XY}(x,y)$. What is the density of their sum, $X + Y$? Since the density is the derivative of the CDF:

\[
p_{X+Y}(z) = \frac{d}{dz} P(X + Y < z) = \frac{d}{dz} \int_{\{(x,y)\mid x+y<z\}} p_{XY}(x,y) \, dx \, dy
\]

\[
= \frac{d}{dz} \int_{\{(x,y)\mid x<z-y\}} p_{XY}(x,y) \, dx \, dy
\]

\[
= \int_{-\infty}^{\infty} \frac{\partial}{\partial z} \left( \int_{-\infty}^{z-y} p_{XY}(x,y) \, dx \right) \, dy
\]

\[
= \int_{-\infty}^{\infty} p_{XY}(z - y, y) \, dy.
\]

This is as far as we can go without further information about $X, Y$. If $X$ and $Y$ are independent, the joint PDF factorizes into a product, $p_{XY}(z - y, y) =$
\( p_X(z - y) \cdot p_Y(y) \), and the last operation becomes a \textit{convolution}:

\[
p_{X+Y}(z) = \int_{-\infty}^{\infty} p_X(z - y) \cdot p_Y(y) dy.
\]

Thus, the PDF of \( Z = X + Y \) is the convolution of the PDFs of \( X \) and \( Y \), whenever \( X \) and \( Y \) are statistically independent.

2.15.3. Distribution of \( \sqrt{X^2 + Y^2} \). Suppose that \( X, Y \sim N(0, \sigma^2) \) (zero-mean Gaussians) are independent rv's. Consider the transformation to polar coordinates:

\[
R = \sqrt{X^2 + Y^2}, \quad \Theta = \tan^{-1}(Y/X).
\]

The inverse transformation is:

\[
x = r \cos \theta, \quad y = r \sin \theta.
\]

What is the distribution of \( \Theta \) and \( R \)? Let us do \( R \). The CDF of \( R \) is found by writing:

\[
P(R < r) = \int_{\{(x,y)|\sqrt{x^2+y^2} < r\}} \frac{1}{2\pi\sigma^2} e^{-(x^2+y^2)/(2\sigma^2)} dx dy.
\]

It will be convenient to transform to polar coordinates. The Jacobian of the transformation yields the new area element:

\[
dx \, dy = \left| \frac{\partial(x, y)}{\partial(r, \theta)} \right| dr \, d\theta,
\]

where

\[
\frac{\partial(x, y)}{\partial(r, \theta)} = \begin{vmatrix} \frac{\partial x}{\partial r} & \frac{\partial x}{\partial \theta} \\ \frac{\partial y}{\partial r} & \frac{\partial y}{\partial \theta} \end{vmatrix} = \begin{vmatrix} \cos \theta & -r \sin \theta \\ \sin \theta & r \cos \theta \end{vmatrix} = r.
\]

Then,

\[
P(R < r) = \frac{1}{2\pi\sigma^2} \int_0^{2\pi} d\theta \int_0^r dr e^{-r^2/(2\sigma^2)} \cdot r = \frac{1}{\sigma^2} \int_0^r dr e^{-r^2/(2\sigma^2)} \cdot r.
\]

The corresponding PDF is obtained by differentiating with respect to \( r \):

\[
p_R(r) = \frac{r}{\sigma^2} e^{-r^2/(2\sigma^2)}.
\]

We have recovered the Rayleigh distribution, by constructing the random variable \( R = \sqrt{X^2 + Y^2} \), where \( X, Y \sim N(0, \sigma^2) \).

The derivation of the distribution for \( \Theta \) is trivial. Recall that \( \Theta = \tan^{-1}(Y/X) \).

Then,

\[
P(\Theta < \theta) = \int_{\{(x,y)|\tan^{-1}(y/x) < \theta\}} \frac{1}{2\pi\sigma^2} e^{-(x^2+y^2)/(2\sigma^2)} dx dy.
\]
2.15. Calculating Probabilities: Two Variables

Transformation to polar coordinates gives:

\[ P(\Theta < \theta) = \int_0^\theta d\theta \int_0^\infty dr \frac{1}{2\pi\sigma^2} e^{-r^2/(2\sigma^2)} r. \]

The integral over \( r \) can be solved with the substitution \( w = r^2/(2\sigma^2), \, dw = r\,dr/\sigma^2 \). Thus, our CDF is:

\[ P(\Theta < \theta) = \frac{1}{2\pi} \int_0^\theta d\theta = \frac{\theta}{2\pi}, \]

where \( \theta \in [0, 2\pi] \). The PDF is that of a uniform distribution:

\[
\begin{align*}
    p_\Theta(\theta) &= \frac{1}{2\pi},
\end{align*}
\]

with \( \theta \in [0, 2\pi] \). Thus, \( R \) is Rayleigh whereas \( \Theta \) is uniform. This can be seen in Fig. 2.5, which is a scatter plot of the pairs \((X, Y)\), where \( X, Y \sim N(0, 1) \).

This plot was generated in MATLAB as follows:

\[
\begin{align*}
    R &= \text{sqrt}(X.^2+Y.^2); \quad \text{theta} = \text{atan}(Y./X); \\
    \text{figure; plot}(X,Y,'.b'); \quad \text{title}(\text{'(X,Y)~N(0,1)'}) \\
    \text{set(gca,'fontsize',16)}; \\
    \text{title('R=(X^2+Y^2)^{1/2}');}
\end{align*}
\]

The distributions of \( R \) and \( \Theta \) can be plotted by taking the pairs \((X, Y)\) and generating \( R, \Theta \). Histograms of \( R \) and \( \Theta \) are shown in Fig. 2.6. It can be seen that \( R \) is Rayleigh and \( \Theta \) is uniform. These plots were generated in MATLAB using the following commands:

\[
\begin{align*}
    R &= \text{sqrt}(X.^2+Y.^2); \quad \text{theta} = \text{atan}(Y./X); \\
    \text{figure; hist}(R,50); \\
    \text{set(gca,'fontsize',16)}; \\
    \text{title('R=(X^2+Y^2)^{1/2}')};
\end{align*}
\]
2. Probability

$$\text{>> figure; hist(theta+pi/2,50); set(gca,'fontsize',16);}$$

$$\text{>> title(''\theta=tan^{-1}\{Y/X\}'');}$$

Figure 2.6. Histograms of $R$ and $\Theta$, as generated from a sequence of random variables $\{(X_i, Y_i)\}_{i=1}^{1000}$, where $X_i, Y_i \sim N(0, 1)$ are all independent.

2.16. Several Variables

Suppose we have several random variables $X_1, X_2, \ldots, X_n$. Probabilistic expressions involving these rv’s can be evaluated if we know the joint distribution:

$$\mathbb{P}(X_1 < b_1, \ldots, X_n < b_n) = \int_{-\infty}^{b_1} dx_1 \cdots \int_{-\infty}^{b_n} dx_n p_{X_1,\ldots,X_n}(x_1, \ldots, x_n),$$

where $p_{X_1,\ldots,X_n}(x_1, \ldots, x_n)$ is the joint PDF of $X_1, \ldots, X_n$ and $\mathbb{P}(X_1 < b_1, \ldots, X_n < b_n)$ is the joint CDF.

We can also ask about the probability of the following event:

$$\{X_1 \in B_1, X_2 \in B_2, \ldots, X_n \in B_n\}.$$

Then, using the joint PDF:

$$\mathbb{P}\{X_1 \in B_1, \ldots, X_n \in B_n\} = \int_{B_1} dx_1 \cdots \int_{B_n} dx_n p_{X_1,\ldots,X_n}(x_1, \ldots, x_n).$$

As in the single-variable case, we can write probabilities over intervals in terms of the CDFs. For example, let $X, Y$ be two random variables and let
2.17. Additional Properties of rv’s

\( B_1 = (a_1, a_2) \) and \( B_2 = (b_1, b_2) \). Then,
\[
\mathbb{P}(X \in B_1, Y \in B_2) = \mathbb{P}(a_1 < X < a_2, b_1 < Y < b_2)
\]
\[
= \int_{a_1}^{a_2} \int_{b_1}^{b_2} dx \, dy \, p_{XY}(x, y)
\]
\[
= \int_{a_1}^{a_2} dx \left( \frac{\partial}{\partial x} \mathbb{P}(X < x, Y < b_2) - \frac{\partial}{\partial x} \mathbb{P}(X < x, Y < b_1) \right)
\]
\[
= \mathbb{P}(X < a_2, Y < b_2) - \mathbb{P}(X < a_1, Y < b_2) \quad - \mathbb{P}(X < a_2, Y < b_1) + \mathbb{P}(X < a_1, Y < b_1).
\]
We have made use of Eq. (2.7), i.e.
\[
p_{XY}(x, y) = \frac{\partial^2}{\partial x \partial y} \mathbb{P}(X < x, Y < y),
\]
and invoked the fundamental theorem of calculus (twice).

2.17. Additional Properties of rv’s

2.17.1. Linearity of the Expectation Value. Let \( X \) and \( Y \) be rv’s and \( a, b \) constants. From the linearity of the expectation value operator:
\[
\mathbb{E}(aX + bY) = \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} (ax + by)p_{XY}(x, y)dx \, dy
\]
\[
= a \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} xp_{XY}(x, y)dx \, dy + b \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} yp_{XY}(x, y)dx \, dy
\]
\[
= a \mathbb{E}(X) + b \mathbb{E}(Y),
\]
where \( p_{XY}(x, y) \) is the joint probability density of \( x \) and \( y \).\(^7\) This can be generalized to any number of random variables, e.g. for \( X = X_1 + X_2 + \cdots + X_n \) it also follows that
\[
\mathbb{E}(X) = \mathbb{E}(X_1) + \mathbb{E}(X_2) + \cdots + \mathbb{E}(X_n).
\]
It is also trivial to show that the same result holds in the case of discrete random variables.

2.17.2. Properties of the Variance. From the definition of the variance of a rv \( X \) (let \( \mu_X \equiv \mathbb{E}(X) \)),
\[
var(X) \equiv \mathbb{E}(X - \mu_X)^2 = \mathbb{E}(X^2) - 2\mathbb{E}(X\mu_X) + \left(\mu_X\right)^2 = \mathbb{E}(X^2) - \mu_X^2
\]
we deduce that
\[
var(aX) = a^2 var(X).
\]
\(^7\)Note: while the exact form of \( p_{XY}(x, y) \) may be unknown, its knowledge was not required to demonstrate linearity.
Example: Let \( v \) be the velocity of a particle and \( t \) the time variable. If both are statistically independent random variables, the mean displacement factors as a product of means, \( \mathbb{E}(vt) = \mathbb{E}v \cdot \mathbb{E}t \), whereas its variance is

\[
\text{var}(vt) = \mathbb{E}((vt)^2) - (\mathbb{E}vt)^2 \\
= \mathbb{E}(v^2t^2) - (\mathbb{E}v)^2(\mathbb{E}t)^2 \\
= \mathbb{E}(v^2) \cdot \mathbb{E}(t^2) - (\mathbb{E}v)^2(\mathbb{E}t)^2.
\]

Thus, by statistically independence, we can express the mean and variance of the displacement \( x = vt \) in terms of the mean and variance of \( v \) and \( t \).

### 2.18. Covariance

The covariance of \( X \) and \( Y \) is defined as (let \( \mu_X \equiv \mathbb{E}X \) and \( \mu_Y \equiv \mathbb{E}Y \)):

\[
\text{cov}(X,Y) \equiv \mathbb{E}(X - \mu_X)(Y - \mu_Y) = \mathbb{E}(XY) - \mu_X \cdot \mu_Y.
\]

We note that the covariance of two independent random variables is zero since \( \mathbb{E}(XY) = \mathbb{E}X \cdot \mathbb{E}Y = \mu_X \cdot \mu_Y \). Using the covariance, we may write the variance of the sum \( X + Y \) as

\[
\text{var}(X + Y) = \mathbb{E}(X + Y - \mathbb{E}(X + Y))^2 \\
= \mathbb{E}(X - \mu_X)^2 + \mathbb{E}(Y - \mu_Y)^2 + 2\mathbb{E}(X - \mu_X)(Y - \mu_Y) \\
= \text{var}(X) + \text{var}(Y) + 2\text{cov}(X,Y).
\]

If \( X \) and \( Y \) are statistically independent, \( \text{cov}(X,Y) = 0 \), and \( \text{var}(X + Y) = \text{var}(X) + \text{var}(Y) \), i.e. the error in \( X + Y \) is related to the errors in \( X \) and \( Y \). In other words, let \( Z = X + Y \), then:

\[
\sigma_Z = \sqrt{\sigma_X^2 + \sigma_Y^2}.
\]

### 2.19. Correlation Coefficient

A concept that is related to the covariance is the correlation coefficient:

\[
\rho(X,Y) = \frac{\text{cov}(X,Y)}{\sigma_X \cdot \sigma_Y},
\]

where \( \sigma_X = \sqrt{\text{var}(X)} \) and \( \sigma_Y = \sqrt{\text{var}(Y)} \). \( \rho \) takes values in the range \(-1 \leq \rho(X,Y) \leq 1 \). The correlation coefficient is a measure of linear dependence between \( X \) and \( Y \). It is more useful than the covariance in the sense that \( \rho \) is a dimensionless quantity which is normalized to the magnitude of \( X \) and \( Y \). A value of \( \rho = 1 \) means that \( X \) and \( Y \) are correlated. A value of \( \rho = -1 \) means they are anti-correlated. A value of \( \rho = 0 \) means they are uncorrelated. Please note that if \( X \) and \( Y \) are independent then \( \rho = 0 \).
However, the converse is not true. That is, $\rho = 0$ does not always imply that $X$ and $Y$ are independent. The reader can easily check\(^8\) that if $Y = aX + b$ ($a, b$ constants) we have $\text{cov}(X, Y) = a \cdot \text{var}(X)$ and

$$
\rho(X, Y) = \frac{a}{|a|} = \begin{cases} 
1 & a > 0 \\
-1 & a < 0
\end{cases}.
$$

Thus, the correlation coefficient is a measure of linear dependence. In this example, $\rho = 1$ if $Y = aX + b$ and $a > 0$ ($X$ and $Y$ are correlated), whereas $\rho = -1$ if $a < 0$ ($X$ and $Y$ are anti-correlated). This result is independent of the magnitude of $a$; it only depends on its sign. For example, $Y = 0.00001 \cdot X$ and $Y = 10000 \cdot X$ both give $\rho = 1$.

Q: Can you find examples of random variables $X$ and $Y$ where $\rho$ is not equal to $-1$, 0 or 1 but some intermediate value (say 0.5)? What is the meaning of a correlation coefficient that is not equal to 0 or 1?

2.20. Sample Correlation Coefficient

Let $X, Y$ be a random variables with mean $\mu_X$ and $\mu_Y$, respectively. Let $x_1, x_2, \ldots, x_n$ be measurements of $X$. Similarly for $Y$. The correlation coefficient can be estimated from experimental data, $x_1, x_2, \ldots, x_n, y_1, y_2, \ldots, y_n$, using the sample correlation coefficient:

$$
r_{X,Y} = \frac{\frac{1}{n-1} \sum_{i=1}^{n} (x_i - \mu_X) \cdot (y_i - \mu_Y)}{\sqrt{\frac{1}{n-1} \sum_{i=1}^{n} (x_i - \mu_X)^2} \cdot \sqrt{\frac{1}{n-1} \sum_{i=1}^{n} (y_i - \mu_Y)^2}}
$$

where $n$ is the number of data points and $\mu_X$ is the sample mean

$$
\mu_X = \frac{1}{n} \sum_{i=1}^{n} x_i,
$$

and similarly for $\mu_Y$. They are uncorrelated if $r_{X,Y} = 0$. You can also check for possible correlation between $X$ and $Y$ using a scatter plot. This is done by plotting the set of ordered pairs $\{(x_i, y_i)\}$ as points on the same graph.

2.21. Uncorrelated but not Independent

Let $X$ and $Y$ be random variables related by $Y = X^2$. Let $\mu_X$ be the mean of $X$, $\mu_{X^2}$ its second moment, etc. Clearly, these random variables are not

\(^8\)Start by the numerator: $\text{cov}(X, Y) = \text{cov}(X, aX + b) = \mathbb{E}(X - \mu_X)(aX + b - \mu_{aX+b})$, but since $\mu_{aX+b} = a\mu_X + b$, this reduces to $\text{cov}(X, aX + b) = \mathbb{E}(X - \mu_X)(aX - a\mu_X) = a\mathbb{E}(X - \mu_X)^2 = a \cdot \text{var}(X)$. 

independent of each other. However,
\[ \text{cov}(X, Y) = \text{cov}(X, X^2) = \mathbb{E}(X - \mu_X)(X^2 - \mu_{X^2}) = \mu_{X^3} - \mu_X \mu_{X^2}. \]

If the distribution of \( X \) is such that \( \mu_{X^3} = \mu_X \mu_{X^2} \) (for example, if the mean and skewness are zero, which is the case for a zero-mean normal distribution), then \( \text{cov}(X, Y) = 0 \) and the random variables \( X \) and \( Y \) are uncorrelated even though they are clearly dependent on each other.

This can easily be illustrated in MATLAB. Let’s create two plots. One for the equation \( Y = X + \eta \) (linear case), where \( X \) and \( \eta \) are independent standard normal rv’s:

```matlab
1  X=randn([1 10000]);
2  Y=X+randn([1 10000]);
```

and one for \( Y = X^2 + \eta \) (quadratic case), where \( X \) and \( \eta \) are independent standard normal rv’s.

```matlab
1  X=randn([1 10000]);
2  Y=X.^2+randn([1 10000]);
```

In each case, a plot is generated by typing `figure;plot(X,Y,'.')` (see Fig. 2.7). The sample correlation coefficient is obtained using the `corrcoef` command, e.g. for the linear case

```matlab
>> corrcoef(X,Y)
ans =
    1.0000    0.7025
    0.7025    1.0000
```
whereas for the quadratic case we get:

```matlab
>> corrcoef(X,Y)
```

```
an =
    1.0000   -0.0089
     -0.0089    1.0000
```

The diagonal elements are 1 (since $X$ is perfectly correlated to $X$, as is $Y$ correlated to $Y$). The off-diagonal elements are the correlation coefficients of $X$ and $Y$. In the linear case, we have strong ($\approx 0.7$) correlation between $X$ and $Y$. We would expect 1 without the additive noise, $\eta$ (you can check this by reducing the amplitude of the additive noise). In the quadratic case, the correlation coefficient is nearly 0, as it should because the Gaussian rv $X$ has zero skewness and zero mean.

### 2.22. General Method for Calculating Probabilities

If you are asked to compute the probability of a random event involving $X$, your first reflex should be to write down an integral of the PDF, $p_X(x)$, over the set of points that represent this event. Recall the formula (2.9) introduced in the previous lecture:

\[ P(X \in A) = \int_{\{x|x \in A\}} p_X(x)dx, \tag{2.9} \]

In two dimensions you do the same thing except that the joint PDF is to be used, e.g.

\[ P(a \leq X \leq b, c \leq Y \leq d) = \int_{\{(x,y)|a \leq x \leq b, c \leq y \leq d\}} p_{XY}(x,y)dx dy = \int_a^b \int_c^d p_{XY}(x,y)dx dy. \]

This is for the specific case where the random even is $\{a \leq X \leq b, c \leq Y \leq d\}$. For a general random event, we integrate over the set of points $(x, y)$ such that the event is true. It is not possible to enumerate all possible events we may encounter. Some examples are:

- $\{X \in A, Y \in B\}$,
- $\{X/Y < a\}$,
- $\{X + Y > b\}$,
- $\{a < \cos(X) \cdot \log(Y) < b\}$, etc.

In each case, we integrate the joint PDF of $X$ and $Y$, $p_{XY}(x, y)$, over the set of points $(x, y)$ that obey the conditions specified in the event.

To summarize the procedure involved when calculating probabilities, there are two main steps. The first step is to write down the right hand side, but
keeping in mind that the random event will need to be expressed in a form suitable for integration. The second step involves writing the integral in a form that can be solved. This sometimes involves a change of variables, if the integration region needs to be simplified.

2.2.2.1. Single-Variable Case. In the first step, we often invoke some algebraic manipulations in order to transform the logical statement \(X \in A\) into a form that allows us to apply the information known to us. Let us revisit the example already covered in Sections 2.8 and 2.7.4. Let \(Y = e^X\) and \(X \sim \mathcal{N}(\mu, \sigma^2)\). You are asked to find the distribution of \(Y\) given the distribution \(X\) (a normal law in the present case). At first sight, you may think that \(\mathbb{P}(Y < y)\) cannot be easily calculated because you are not given the distribution of \(Y\). However, the distribution of \(X\) is provided. So your goal is to transform the statement \(Y < y\) into one that involves \(X\) instead.

First we note that the following three probabilities are all equal:

\[
\mathbb{P}(Y < y) = \mathbb{P}(e^X < y) = \mathbb{P}(X < \log y).
\]

The quantity \(\mathbb{P}(X < \log y)\) is something we can calculate because the distribution of \(X\) is known. Application of formula (2.9) gives:

\[
\mathbb{P}(X < \log y) = \int_{-\infty}^{\log y} \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{(x-\mu)^2}{2\sigma^2}} \, dx.
\]

Therefore,

\[
\mathbb{P}(Y < y) = \int_{-\infty}^{\log y} \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{(x-\mu)^2}{2\sigma^2}} \, dx = \text{Erf}(\log y)_{\mu,\sigma} = \Phi \left( \frac{\log y - \mu}{\sigma} \right),
\]

where \(\Phi(z)\) is the normalized error function (see Sections 2.8 and 2.7.4). We cannot simplify this any further, because the definite integral does not have a result in closed form. Instead we must calculate its value numerically. For example, you can use the MATLAB command \texttt{normcdf} to get its value.

2.2.2.2. Two Variables Case. Another example is \(Y = U/V\) where \(U\) and \(V\) are independent standard normal variables, i.e \(U \sim \mathcal{N}(0, 1)\) and \(V \sim \mathcal{N}(0, 1)\). The CDF of \(Y\) is:

\[
\mathbb{P}(Y < a) = \mathbb{P}(U/V < a) = \int \int_{\{(u,v)|u/v<a\}} p_U(u)p_V(v) \, du \, dv
\]

Effecting a change of variables \(Y = U/V, Z = V\) (inverse: \(V = Z, U = ZY\)) under the integral sign and using the Jacobian of the transformation:

\[
dudv = \left| \frac{\partial(u,v)}{\partial(y,z)} \right| \, dydz, \quad \text{where} \quad \left| \frac{\partial(u,v)}{\partial(y,z)} \right| = \begin{vmatrix} x & y \\ x & z \end{vmatrix} = z
\]

\[
\begin{vmatrix} \frac{\partial_y u}{\partial_y v} & \frac{\partial_z u}{\partial_z v} \\ \frac{\partial_y v}{\partial_y z} & \frac{\partial_z v}{\partial_z z} \end{vmatrix} = \begin{vmatrix} x & y \\ x & z \end{vmatrix} = z
\]

\[
\mathbb{P}(Y < a) = \int \int_{\{(u,v)|u/v<a\}} p_U(u)p_V(v) \, du \, dv
\]

is:
where $\| \cdot \|$ denotes “matrix determinant”, whereas $| \cdot |$ denotes absolute value. Then,

\[
P(Y < a) = \int \int_{\{(y,z) | y < a\}} p_U(yz)p_V(z)|z|dydz
\]

\[
= \frac{1}{2\pi} \int_{-\infty}^{\infty} \left\{ \int_{-\infty}^{a} e^{-\frac{1}{2}y^2} e^{-\frac{1}{2}z^2}|z|dy \right\} dz.
\]

Using the change of variables $w = \frac{1}{2}z^2$, $dw = zdz$ and replacing $\int_{-\infty}^{\infty} dz$ by $2 \int_{0}^{\infty} dz$ (this replacement is allowed since its integrand is positive):

\[
P(Y < a) = \frac{1}{\pi} \int_{-\infty}^{a} \left\{ \int_{0}^{\infty} e^{-\frac{1}{2}z^2(1+y^2)}zdz \right\} dy
\]

\[
= \frac{1}{\pi} \int_{-\infty}^{a} \left\{ \int_{0}^{\infty} e^{-w(1+y^2)}dw \right\} dy = \int_{-\infty}^{a} \frac{1}{\pi} \frac{dy}{1+y^2}.
\]

This is known as the Lorentzian (or Cauchy) distribution. The PDF of the Lorentzian distribution is obtained by differentiating with respect to $a$:

\[
p_Y(y) = \frac{1}{\pi} \frac{1}{(1+y^2)}.
\]

In another example we can ask what is the probability that a rv $X$ takes on exactly the value $x$:

\[
P(X = x) = \lim_{dx \to 0} P(x < X \leq x + dx) = \lim_{dx \to 0} \int_{x}^{x+dx} p(x)dx = \lim_{dx \to 0} p(x)dx = 0
\]

provided that $p(x)$ is continuous. If $p(x)$ is discontinuous at $x$, this result is not necessarily zero. In this course, we will not be dealing with discontinuous probability functions.

### 2.23. Probability of Mutually Exclusive Random Events

If random events $A_1$, $A_2$ and $A_n$ are disjoint sets, i.e. $A_i \cap A_j = \emptyset$, then the probability of any of the $A_i$ events is a sum of probabilities:

(2.10) \[ \mathbb{P}(A_1 \cup A_2 \cup \cdots \cup A_n) = \mathbb{P}(A_1) + \mathbb{P}(A_2) + \cdots + \mathbb{P}(A_n). \]

Such a set of random events is called mutually exclusive events. The “union” $A_1 \cup A_2 \cup \cdots \cup A_n$ of random events is equivalent to a “logical OR” operation, i.e.

\[ \mathbb{P}(A_1 \cup A_2 \cup \cdots \cup A_n) = \mathbb{P}(A_1 \text{ or } A_2 \text{ or } \ldots \text{ or } A_n). \]

When we make the statement “12 or fewer counts”, we mean to say:

12 or fewer counts = 1 count or 2 counts or ... or 12 counts.
Figure 2.8. Mutually exclusive random events $A$ and $B$.

Figure 2.9. Events that are not mutually exclusive share common outcomes (as represented here by the overlap region).

In other words, let $X$ be a random variable that represents the # of counts (in some given time interval).

$$\{X \leq 12\} = \{X = 1\} \cup \{X = 2\} \cup \{X = 3\} \cup \cdots \cup \{X = 12\}.$$  

Decomposing the event $\{X \leq 12\}$ as a union of mutually exclusive random events, i.e. $\{X = 1\} \cap \{X = 2\} = \emptyset$, etc., offers some advantages when calculating the probabilities of events. It enables us to invoke formula 2.10.

Two mutually exclusive events $A$ and $B$ have no overlap can be represented as shown in Fig. 2.8. What should we do if the random events are not mutually exclusive? For simplicity, consider only 2 events, $A$ and $B$. Mutual exclusivity means that $A \cap B = \emptyset$. If the intersection is nonzero, then we have the situation illustrated in the Venn diagram (Fig. 2.9).

In this case, we should avoid overcounting by subtracting the intersection:

$$\mathbb{P}(A \cup B) = \mathbb{P}(A) + \mathbb{P}(B) - \mathbb{P}(A \cap B).$$

Let us look at the example of rolling a die. Let $X$ be the numerical result of the die roll (# appearing on the top face). Define two random events, $A$ and $B$, as follows: $A = \{X \leq 3\}$ and $B = \{X \text{ is odd}\}$. It is helpful to decompose these random events into a union of mutually exclusive “elementary events”:

$$A = \{X \leq 3\} = \{X = 1\} \cup \{X = 2\} \cup \{X = 3\},$$

and

$$B = \{X \text{ is odd}\} = \{X = 1\} \cup \{X = 3\} \cup \{X = 5\}.$$
The union of $A$ and $B$ is:

$$A \cup B = \{X = 1\} \cup \{X = 2\} \cup \{X = 3\} \cup \{X = 5\}.$$

The intersection of $A$ and $B$ is:

$$A \cap B = \{X = 1\} \cup \{X = 3\}.$$

If the die is unbiased, i.e. $P(X = x_i) = 1/6$ for $x_i \in \{1, 2, \ldots, 6\}$, then $P(A \cup B) = 2/3$, $P(A) = 1/2$, $P(B) = 1/2$ and $P(A \cap B) = 1/3$. This verifies the formula $P(A \cup B) = P(A) + P(B) - P(A \cap B)$ for this particular choice of random events.

### 2.24. Discrete Random Variables

A discrete random variable $X$ takes values in a discrete set $\{x_i\}_{i=1}^N$, where $N$ is the number of possible values $X$ can take and $x_i$ are the possible values. The word discrete refers to the “state space”, which is discrete and contains $N$ elements.

#### 2.24.1. Properties of Discrete Random Variables

The random variable is defined by the probability distribution $\{p_i\}$ (also known as the “probability mass function” or PMF), where $p_i \geq 0$ for all $i$. The normalization condition is

$$\sum_{i=1}^N p_i = 1$$

$p_i$: probability that random variable $X$ takes the value $x_i$.

$N$: can be finite or infinite; in any case, the $\{p_i\}$ must sum to 1.

We note that from the definition of $p_i$ as the probability that $X$ takes the discrete value $x_i$, and the fact that the events $\{X = x_i\}$ are mutually exclusive random events, it follows that

$$P(X \leq x_j) = P(\{X = x_1\} \cup \{X = x_2\} \cup \cdots \cup \{X = x_j\}) = \sum_{i=1}^j P(\{X = x_i\}) = \sum_{i=1}^j p_i.$$ 

The mean value of $X$ is

$$\mathbb{E}X = \sum_{i=1}^N p_i x_i.$$ 

The $n$-th moment of $X$ is

$$\mathbb{E}(X^n) = \sum_{i=1}^N p_i x_i^n.$$
The mean or mathematical expectation of $f(X)$ is
\[ \mathbb{E}(f(X)) \equiv \sum_{i=1}^{N} p_i f(x_i). \]

The variance of $X$ is (let $\mu_X = \mathbb{E}X$)
\[ \sigma^2 = \text{Var}(X) = \mathbb{E}(X - \mu_X)^2 = \sum_{i=1}^{N} p_i (x_i - \mu_X)^2. \]

The variance is also equal to $\mathbb{E}(X^2) - (\mu_X)^2$. When calculating averages of functions of random variables, we proceed by replacing the random variable $X$ by its value $x_i$, multiplying the expression by $p_i$ and summing over all $i$. For example,
\[ \mathbb{E}(\exp(-X)) = \sum_{i} p_i \exp(-x_i), \quad \mathbb{E}(g(X)) \equiv \sum_{i} p_i g(x_i). \]

This is analogous to the continuous case covered in the previous lecture where \( \int p(x) \) replaces \( \sum_i p_i \):
\[ \mathbb{E}(\exp(-X)) = \int_{-\infty}^{\infty} p(x) \exp(-x)dx, \quad \mathbb{E}g(X) = \int_{-\infty}^{\infty} p(x)g(x)dx. \]

When we deal with two discrete random variables $X$ and $Y$, the joint probability can be written
\[ p_{ij} \equiv P(X = x_i \cap Y = y_j) \]
where $P(X = x_i \cap Y = y_j)$ denotes the probability of $X$ taking the value $x_i$ and $Y$ taking the value $y_j$. The $p_{ij}$ are normalized to 1:
\[ \sum_{i,j} p_{ij} = 1. \]

The average of a function $g(X,Y)$ is:
\[ \mathbb{E}g(X,Y) = \sum_{ij} p_{ij} g(x_i,y_j). \]

As before, if $X$ and $Y$ are statistically independent, the mean of $XY$ equals the product of the means of $X$ and $Y$:
\[ \mathbb{E}(XY) = \sum_{ij} x_i y_j P(X = x_i \cap Y = y_j) \]
\[ = \sum_{i} x_i P(X = x_i) \sum_{j} y_j P(Y = y_j) = \mathbb{E}X \cdot \mathbb{E}Y, \]
and similarly, we have:
\[ \mathbb{E}(X^n Y^m) = \mathbb{E}(X^n) \cdot \mathbb{E}(Y^m). \]
2.24. Discrete Random Variables

A consequence of this result is

$$E(f(X)g(Y)) = E(f(X)) \cdot E(g(Y)),$$

because sufficiently nice functions \( f, g \) can be expanded as a power series, enabling us to apply the result \( E(X^nY^m) = E(X^n) \cdot E(Y^m) \) to each term of the expansion.

2.24.2. Poisson Distribution. The Poisson distribution is a discrete probability distribution which is frequently used to describe counts of rare events. The main assumptions are:

- The events counted are rare events.
- All events are statistically independent.
- Average count rate does not change over time.

The typical application of this distribution is radioactive counting (for example, with a Geiger counter), where the average count \( \bar{\pi} \) in a given time is given by the formula:

$$n = \lambda \tau$$

where \( \lambda \) is average count rate and \( \tau \) is time interval. For example: \( \lambda = 1.5 \) s\(^{-1} \), \( \tau = 10 \) s gives \( \bar{\pi} = 15 \). \( \bar{\pi} \) does not have to be an integer number.

2.24.2.1. Probability Mass Function. The probability distribution is completely described by a single parameter, \( \bar{\pi} \):

$$P(n; \bar{\pi}) = \frac{e^{-\bar{\pi}}\bar{\pi}^n}{n!}, \quad n = 0, 1, 2, \ldots$$

In the usual interpretation, \( P(0; \bar{\pi}) \) gives the probability of observing 0 counts in a time interval \( \tau \), \( P(1; \bar{\pi}) \) gives the probability of observing 1 count, etc. It is easy to check that this PMF is normalized to 1:

$$\sum_{n=0}^{\infty} P(n; \bar{\pi}) = \sum_{n=0}^{\infty} \frac{e^{-\bar{\pi}}\bar{\pi}^n}{n!} = e^{-\bar{\pi}} \sum_{n=0}^{\infty} \frac{\bar{\pi}^n}{n!} = e^{-\bar{\pi}} e^{\bar{\pi}} = 1,$$

since the Taylor expansion of \( \exp(x) \) is \( \sum_{k=0}^{\infty} x^k / k! \). Figure 2.10 shows plots of the Poisson distribution for \( \bar{\pi} = 3 \) and \( \bar{\pi} = 10 \). Notice that the distribution looks more like a Gaussian at large \( \bar{\pi} \).

These plots were generated in MATLAB using the commands:

```matlab
1 Nbar=2;N=0:20;figure;plot(N,exp(-Nbar)*(Nbar.'^N)./factorial(N),'o');
2 Nbar=5;N=0:20;figure;plot(N,exp(-Nbar)*(Nbar.'^N)./factorial(N),'o');
3 Nbar=10;N=0:20;figure;plot(N,exp(-Nbar)*(Nbar.'^N)./factorial(N),'o');
4 figure;ezplot('exp(-((x-10)^2)/(2*10))',[0,20]);
```

\(^9\)This is in contrast to the Gaussian distribution, whose description requires two parameters: \( \bar{X} \) and \( \sigma^2 \).
Two more properties of the Poisson distribution which you can easily verify are:

$$\text{average of } n = \sum P(n; \lambda) n = \lambda, \quad \text{var}(n) = \sigma^2 = \lambda = \lambda \tau.$$ 

Thus, the mean and variance are both equal to \( \lambda \). The standard deviation, \( \sigma = \sqrt{\lambda} \), gives the error in the measurement. For the mean, the proof is trivial and left as an exercise. For the variance, the proof is easy but requires more steps.\(^\text{10}\)

2.24.2.2. Error Bars of a Measurement. If the experiment yields a mean count of \( n \), the best estimate of the error\(^\text{11}\) in this quantity is \( \sqrt{n} \). We report this measurement as 

$$n \pm \sqrt{n}.$$ 

While it may seem that the error grows with \( n \), the fractional uncertainty actually decreases with \( n \):

$$\frac{\delta n}{n} = \frac{\sqrt{n}}{n} = \frac{1}{\sqrt{n}},$$

i.e., larger \( n \) result in smaller fractional uncertainty.

2.24.2.3. Poisson Counts. A bridge cannot hold too many cars at once due to the potential for structural damage. A particular bridge is designed to hold less than 13 cars (safe level) per time interval (1 min.). In a random sample, the total number of cars recorded crossing the bridge in 10 hours was 1980.

\(^{10}\)A proof can be found at: [https://proofwiki.org/wiki/Variance_of_Poisson_Distribution](https://proofwiki.org/wiki/Variance_of_Poisson_Distribution)

\(^{11}\)Sometimes, all we have is 1 count. While this may not be the mean count, it is all that we have. The best we can do in this case is report \( n \pm \sqrt{n} \).
Q: What is the average number of cars crossing per minute and its error?  
A:  \( \lambda = \frac{1980}{10 \times 60} = 3.30 \text{ cars/min.} \) 

\[ \delta \lambda = \sqrt{\frac{1980}{10 \times 60}} = 0.07 \text{ cars/min.} \]

Q: What is the probability that during a random one-minute interval 13 cars will be observed crossing?  
A:  \( \pi = \lambda \tau = 3.30, \) with \( n = 13 \) 

\[ \mathbb{P}(n = 13; 3.3) = \frac{e^{-3.3}3.3^{13}}{13!} = 3.3 \times 10^{-5} \]

Q: What is the probability that the bridge will fail (due to too many cars crossing)?  
A:  

\[ \mathbb{P}(13 \text{ or more cars}) = 1 - \mathbb{P}(12 \text{ or fewer cars}) \]

\[ = 1 - \{ \mathbb{P}(0; 3.3) + \mathbb{P}(1; 3.3) + \cdots + \mathbb{P}(12; 3.3) \} \]

\[ = 4.2 \times 10^{-5}. \]

During 1 minute of observation, this is the probability that the bridge will fail.

2.24.2. Poisson Distribution in the Limit of Large Means. For large means, the Poisson distribution converges to a Gaussian distribution:

\[ \frac{e^{-\bar{n}}\bar{n}^n}{n!} \approx \frac{1}{\sqrt{2\pi \mu_X}} \exp \left[ -\frac{(x - \mu_X)^2}{2\mu_X} \right] \]

where:

- \( n \) (discrete) \( \to x \) (continuous)
- \( \bar{n} \to \mu_X \) (variance, mean)
- \( \sqrt{n} \to \sqrt{\mu_X} \) (standard deviation)

The proof makes use of Stirling’s approximation

\[ \log n! \approx n \log n - n + O(\log n), \]

and

\[ \left| \frac{n - \bar{n}}{\bar{n}} \right| \ll 1. \]

These two conditions (Stirling approx. and \( \bar{n} \) close to \( n \)) imply that our proof is valid in the limit of large means. If the mean is not large, the Stirling approximation cannot be used.
\[
\frac{e^{-\pi n/n}}{n!} = \exp \{-\pi - \log n! + n \log \pi\}
\]
\[
= \exp \{-\pi - n \log n + n + n \log \pi\}
\]
\[
= \exp \{(n - \pi) + n \log(\pi/n)\}
\]
\[
= \exp \left\{ (n - \pi) + n \log \left[ 1 + \left( \frac{\pi - n}{n} \right) \right] \right\}
\]
\[
\approx \exp \left\{ -\frac{(\pi - n)^2}{2n} \right\} \approx \exp \left\{ -\frac{(\pi - n)^2}{2\pi} \right\}
\]

The first step was to invoke the Stirling’s approximation, \(\log n! \approx n \log n - n\). The second step was to expand about mean \((\pi)\) for large \(\pi\). Then we Taylor expanded \(\log(1 + x) \approx x - x^2/2 + O(x^3)\). In the last step we have used the approximation \(n \approx \pi\) near the mean for the denominator in the argument of the exp.

The prefactor \(1/\sqrt{2\pi n}\) could have been recovered had we used the slightly more accurate form of the Stirling’s formula:
\[
n! \approx \sqrt{2\pi n} \left( \frac{n}{e} \right)^n,
\]
and of course, followed by the application of \(|n - \pi|/\pi \ll 1\) to justify replacing \(\sqrt{2\pi n}\) by \(\sqrt{2\pi \pi}\). See Problem 32.

2.24.3. Statistical Independence (Discrete Case). The notion of statistical independence is the same as in the continuous case. Namely, the joint probability factors as a product:
\[
(2.11) \quad p_{ij}^{XY} = p_i^X \cdot p_j^Y, \text{ for all } i, j
\]
where \(p^{XY}\) is the joint PMF for \(X\) and \(Y\). \(p_i^X\) is the marginal PMF:
\[
p_i^X = \sum_j p_{ij}^{XY}.
\]
Likewise, \(p_j^Y\) is also a marginal PMF:
\[
p_j^Y = \sum_i p_{ij}^{XY}.
\]

2.24.4. Example 1. Consider a random experiment that involves rolling a die
\[
X \in \{1, 2, 3, 4, 5, 6\}
\]
and tossing a coin
\[
Y \in \{H, T\}.
\]
You are asked to determine whether or not \( X \) is statistically independent from \( Y \). Intuitively, this should be the case (i.e., why would a coin toss affect the outcome of rolling a die?).

For this random experiment, there are 12 possible “elementary” outcomes:

\[
(X, Y) = (1, H), \quad (X, Y) = (1, T),
\]
\[
(X, Y) = (2, H), \quad (X, Y) = (2, T),
\]
\[
(X, Y) = (3, H), \quad (X, Y) = (3, T),
\]
\[
(X, Y) = (4, H), \quad (X, Y) = (4, T),
\]
\[
(X, Y) = (5, H), \quad (X, Y) = (5, T),
\]
\[
(X, Y) = (6, H), \quad (X, Y) = (6, T).
\]

To get the joint PMF we must repeat this experiment many times and record the results. Suppose that we repeat the experiment 10,000 times and count the number of times each elementary outcome occurs. Let’s do this in MATLAB:

```matlab
>> X=randi([1 2],[1 10000]);
>> Y=randi([1 6],[1 10000]);
```

We then plot a 2D histogram (see Fig. 2.11):

```matlab
>> figure; h=histogram2(X,Y)
```

h =

```
Histogram2 with properties:

Data: [10000x2 double]
Values: [2x6 double]
NumBins: [2 6]
XBinEdges: [0.5000 1.5000 2.5000]
YBinEdges: [1x7 double]
BinWidth: [1 1]
Normalization: 'count'
FaceColor: 'auto'
EdgeColor: [0.1500 0.1500 0.1500]
```

The histogram is an approximation to the PMF. As you can see, this distribution is uniform. The values used to plot the histogram are:

```matlab
>> h.Values
```
Dividing by the number of experiments performed (10,000):

```matlab
>> h.Values/10000
```

```
ans =
0.0818 0.0816 0.0807 0.0827 0.0870 0.0831
0.0861 0.0847 0.0824 0.0837 0.0811 0.0851
```

gives an approximation to the joint PMF. Each entry is approximately equal to 1/12. A PMF with entries equal to 1/12 can be represented as a matrix (see Table 2.2). The marginal PMFs for $X$ and $Y$ are obtained by summing along rows and columns, respectively. You can check that for the data shown in Table 2.2), Eq. (2.11) holds for all $i,j$. Therefore, $X$ and $Y$ are statistically independent.

We also know that statistical independence implies the variables are uncorrelated. Let’s check this by computing the correlation coefficient:

```matlab
>> corrcoef(X,Y)
```

```
ans =
1.0000 -0.0115
-0.0115 1.0000
```
The MATLAB command `corrcoef` computes the matrix of correlation coefficients, 
\[
\begin{bmatrix}
\rho(X,X) & \rho(X,Y) \\
\rho(Y,X) & \rho(Y,Y)
\end{bmatrix}
\]. The diagonal elements should be 1 (since $X$ is fully correlated with $X$; same for $Y$) whereas the off-diagonal elements should be zero. Indeed, the off-diagonal elements are two orders of magnitude smaller than 1, indicating the lack of correlation between $X$ and $Y$.

A counter-example illustrating statistical dependence would be the joint PMF shown in Table 2.3, which differs from Table 2.2 only in the second row. Namely, when the result from rolling the die is 2, the coin toss always yields “tails”. (Don’t try too hard to imagine how this can possibly happen in the lab; it is perhaps easier to imagine that a magician is doing the experiment for you.) Because $p_{2,1}^{XY} = 0 \neq p_2^X \cdot p_1^Y = \frac{1}{6} \cdot \frac{1}{2}$, we are unable to prove statistical independence of $X$ and $Y$. 

![Table 2.2. Joint PDF for an experiment that involves rolling a die ($X \in \{1, 2, 3, 4, 5, 6\}$) and tossing a coin ($Y \in \{H, T\}$).](image)

<table>
<thead>
<tr>
<th>X \ Y</th>
<th>heads</th>
<th>tails</th>
<th>sum</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>2</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>3</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>4</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>5</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>6</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>sum</td>
<td>1/2</td>
<td>1/2</td>
<td></td>
</tr>
</tbody>
</table>

![Table 2.3. Joint PDF for an experiment that involves rolling a die ($X \in \{1, 2, 3, 4, 5, 6\}$) and tossing a coin ($Y \in \{H, T\}$). This joint PMF is the same as that of Table 2.2 except for the second row.](image)

<table>
<thead>
<tr>
<th>X \ Y</th>
<th>heads</th>
<th>tails</th>
<th>sum</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>2</td>
<td>0/12</td>
<td>1/6</td>
<td>1/6</td>
</tr>
<tr>
<td>3</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>4</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>5</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>6</td>
<td>1/12</td>
<td>1/12</td>
<td>1/6</td>
</tr>
<tr>
<td>sum</td>
<td>1/2</td>
<td>1/2</td>
<td></td>
</tr>
</tbody>
</table>
2.24.5. Example 2. The joint distribution of the bivariate random variable \((X, Y)\) is given by
\[
p_{XY}(x_i, y_j) = \begin{cases} 
  k \frac{|x_i|}{2y_j} & x_i = -1, 1; y_j = 1, 2, 3, \ldots \ (\text{to infinity}) \\
  0 & \text{otherwise}
\end{cases}
\]
(a) \(k\) is a constant. Find the value of \(k\).
\[
\sum_i \sum_j p_{ij} = k2 \sum_j \frac{1}{2y_j} = k2 \cdot 1 = 2k. \quad k = 1/2.
\]
(b) Find the marginal probability distributions of \(X\) and \(Y\).
\[
p_X(x_i) = \sum_j p_{ij} = \sum_j \frac{1}{2} |x_i| \frac{1}{2y_j} = \frac{1}{2} |x_i|. \quad x_i = -1, 1.
\]
\[
p_Y(y_j) = \sum_i p_{ij} = \frac{1}{2y_j}. \quad y_j = 1, 2, 3, \ldots
\]
(c) Are \(X\) and \(Y\) statistically independent? Forming the product of marginal distributions,
\[
p_X(x_i)p_Y(y_j) = \frac{1}{2} |x_i| \cdot \frac{1}{2y_j} = p_{XY}(x_i, y_j)
\]
Hence \(X\) and \(Y\) are independent.

2.24.6. Example 3. Suppose that we have an amplifier that takes a voltage and amplifies it by a factor of 10\(x\), i.e. \(f(x) = 10x\). Suppose that we feed this amplifier the following input voltages:
\[X = \{2.53, 2.55, 2.45, 2.49, 2.50, 2.52, 2.47, 2.48, 2.56, 2.49\}\]
(a) What is the sample variance at the output of the amplifier? First, calculate the sample variance of \(X\) and then multiply by 100. Then multiply each \(X_i\) by 10 and then calculate the sample variance of the multiplied values.
The sample variance of the \(X\)'s is 0.001249. Multiply this by 100 to get 0.1249. Multiplying each \(X\) by 10 and taking the sample variance we get 0.1249, which is the same as the first method. From this we confirmed the validity of the formula \(\text{var}(aX) = a^2 \text{var}(X)\).
(b) Suppose that we have two random variables, \(X\) and \(Y\) and they are statistically independent. Furthermore, suppose that \(\text{var}(X) = 2.7\) and \(\text{var}(Y) = 2.5\). Compute the value of \(\text{var}(X + Y)\) and \(\text{var}(X - Y)\).
By statistical independence we have
\[
\text{var}(X + Y) = \text{var}(X) + \text{var}(Y) = 2.7 + 2.5 = 5.2.
\]
Then from
\[ \text{var}(aX + bY) = a^2 \text{var}(X) + b^2 \text{var}(Y), \]
with \( a = 1 \) and \( b = -1 \) we have
\[ \text{var}(X-Y) = (1)^2 \text{var}(X)+(-1)^2 \text{var}(Y) = \text{var}(X)+\text{var}(Y) = 2.7+2.5 = 5.2. \]
From this, we conclude that when \( X \) and \( Y \) are statistically independent, \( \text{var}(X+Y) = \text{var}(X-Y) = \text{var}(X)+\text{var}(Y). \)

(c) Given that \( \text{var}(X)=2.7, \text{var}(Y)=2.5 \) and \( \rho(X,Y)=0.9 \) (correlation coefficient), what is \( \text{var}(X+Y) \) and \( \text{var}(X-Y) \)?

From the definition of the correlation coefficient
\[ \rho(X,Y) = \frac{\text{cov}(X,Y)}{\sqrt{\text{var}(X)\text{var}(Y)}} \]
we have \( \text{cov}(X,Y) = \rho(X,Y)\sqrt{\text{var}(X)\text{var}(Y)} \). Then, inserting this into:
\[ \text{var}(X+Y) = \text{var}(X) + \text{var}(Y) + 2\text{cov}(X,Y) \]
we get:
\[ \text{var}(X+Y) = \text{var}(X) + \text{var}(Y) + 2\rho(X,Y)\sqrt{\text{var}(X)\text{var}(Y)} \]
from which we can obtain a numerical value:
\[ \text{var}(X+Y) = 2.7 + 2.5 + 2(0.9)[(2.7)(2.5)]^{1/2} = 9.877. \]

For \( \text{var}(X-Y) \) we have:
\[ \text{var}(X-Y) = \text{var}(X) + \text{var}(-Y) + 2\rho(X,-Y)\sqrt{\text{var}(X)\text{var}(-Y)} \]
and since \( \text{var}(-Y) = (-1)^2 \text{var}(Y) = \text{var}(Y) \), we have:
\[ \text{var}(X-Y) = \text{var}(X) + \text{var}(Y) + 2\rho(X,-Y)\sqrt{\text{var}(X)\text{var}(Y)} \]
Now, using the property \( \text{cov}(aX,bY) = ab \cdot \text{cov}(X,Y) \), we see that
\[ \rho(aX,bY) = \frac{\text{cov}(aX,bY)}{\sqrt{\text{var}(aX)\text{var}(bY)}} = \frac{ab \cdot \text{cov}(X,Y)}{|a||b|\sqrt{\text{var}(X)\text{var}(Y)}} = \frac{ab}{|a||b|}\rho(X,Y). \]
For \( a = 1 \) and \( b = -1, \)
\[ \rho(1 \cdot X,-1 \cdot Y) = \frac{(1) \cdot (-1)}{|1||-1|}\rho(X,Y) = -\rho(X,Y). \]
Hence,
\[ \text{var}(X-Y) = \text{var}(X) + \text{var}(Y) - 2\rho(X,Y)\sqrt{\text{var}(X)\text{var}(Y)} \]
and then we have
\[ \text{var}(X-Y) = 2.7 + 2.5 - 2(0.9)[(2.7)(2.5)]^{1/2} = 0.5235. \]
We conclude that \( \text{var}(X+Y) \) is not equal to \( \text{var}(X-Y) \) when \( X \) and \( Y \) are correlated to some extent.
(d) If \( X, Y \) and \( Z \) are statistically independent and \( \text{var}(X)=1.7 \), \( \text{var}(Y)=2.3 \) and \( \text{var}(Z)=1.4 \). What is \( \text{var}(0.3X + 0.7Y + 0.5Z) \)?

If \( X, Y \) and \( Z \) are statistically independent, then we can write:

\[
\text{var}(aX + bY + cZ) = a^2 \text{var}(X) + b^2 \text{var}(Y) + c^2 \text{var}(Z),
\]

where \( a, b, c \) are constants. Hence,

\[
\text{var}(0.3X + 0.7Y + 0.5Z) = (0.3)^2 1.7 + (0.7)^2 2.3 + (0.5)^2 1.4 = 1.630.
\]

2.25. Signal Averaging Reduces Relative Error

An important concept in experimental science is that of “signal averaging”. This is done to reduce the noise error, or equivalently, to reduce the size of the error bars relative to the measurement. Suppose \( X \) is a random variable and we perform \( n \) measurements of \( X \). The values are:

\[ X_1, X_2, \ldots, X_n. \]

Each of which are statistically independent random variables with mean \( \mu_X \) and finite variance \( \text{var}(X_i) < \infty \). We form the average:

\[ X_{av} = \frac{X_1 + X_2 + \cdots + X_n}{n}. \]

The variance of this sum is

\[
\text{var}(X_{av}) = \frac{\sum_{i=1}^{n} \text{var}(X_i)}{n^2} \propto \frac{1}{n},
\]

where we used the property \( \text{var}(aX) = a^2 \text{var}(X) \). The noise is the square root of the variance \( \sigma = \sqrt{\text{var}(X_{av})} \). Thus, signal averaging reduces the noise from random errors.

We have just derived the formula for standard error: the mean of \( X_{av} = \frac{X_1 + X_2 + \cdots + X_n}{n} \) is \( \mu_X \) (where \( \mu_X = \mathbb{E}X \)) and its standard deviation (error) is \( \sqrt{\text{var}(X)/n} = \sigma_X/\sqrt{n} \).

If the \( X_i \) are iidrv the signal-to-noise ratio (SNR) is defined as:

\[
\text{SNR} = \frac{\text{signal}}{\text{noise}} = \frac{X_{av}}{\sqrt{\text{var}(X_{av})}} \propto \sqrt{n}
\]

Thus, SNR improves with signal averaging.
2.26. Some Theorems on Random Variables

2.26.1. Normal Linear Transform Theorem. The normal linear transform theorem is:

$$\alpha + \beta N(\mu, \sigma^2) = N(\alpha + \beta \mu, \beta^2 \sigma^2).$$

Proof: Let \( Y = \alpha + \beta X \), where \( X \sim N(\mu, \sigma^2) \). Write

$$P(Y < a) = P(\alpha + \beta X < a) = P(X < (a - \alpha)/\beta)$$

and make a substitution of variables \( y = \alpha + \beta x \), \( dy = \beta \, dx \) to get an integral of the form

$$\int_{-\infty}^{a} \frac{1}{\sqrt{2\pi\sigma^2}} e^{-((x-\alpha)/\beta - \mu)^2/2\sigma^2} \, dx, \quad \text{and make a substitution of variables } y = x - \alpha. \quad \text{This is the CDF of a normal rv with mean } \alpha + \beta \mu \text{ and variance } \beta^2 \sigma^2.$$

In the special case \( \mu = 0, \sigma^2 = 1 \) we have:

$$\alpha + \beta N(0, 1) = N(\alpha, \beta^2).$$

2.26.2. Normal Sum Theorem. The normal sum theorem is:

$$N(\mu_1 + \mu_2, \sigma_1^2 + \sigma_2^2) = N_1(\mu_1, \sigma_1^2) + N_2(\mu_2, \sigma_2^2),$$

where on the right-hand-side is the sum of two statistically independent rv’s. In other words, let \( X \sim N(\mu_1, \sigma_1^2) \) and \( Y \sim N(\mu_2, \sigma_2^2) \) be independent rv’s. We are asking what is the distribution of the new random variable \( U = X + Y \). The proof of this involves handling a 2D integral of the joint PDF of \( X \) and \( Y \):

$$P(X + Y < a) = \frac{1}{2\pi\sigma_1\sigma_2} \int_{\{(x,y)|x+y<a\}} e^{-\frac{(x-\mu_1)^2}{2\sigma_1^2}} e^{-\frac{(y-\mu_2)^2}{2\sigma_2^2}} \, dx \, dy.$$
start with the CDF:
\[ P(X + Y < a) = \int_{\{(x,y)|x+y<a\}} \frac{1}{\pi^2} \frac{1}{1+x^2} \frac{1}{1+y^2} \, dx \, dy. \]

Let us effect a change of variables: \( u = x + y \) and \( v = y \). The inverse is \( y = v \) and \( x = u - v \). The area element is:
\[ dx \, dy = \left| \frac{\partial(x,y)}{\partial(u,v)} \right| du \, dv = \left| \det \begin{pmatrix} 1 & 1 \\ 0 & 1 \end{pmatrix} \right| du \, dv = du \, dv. \]

In the new coordinates, the integral is:
\[ P(X + Y < a) = \frac{1}{\pi^2} \int_{-\infty}^{a} du \int_{-\infty}^{\infty} dv \frac{1}{(1+(u-v)^2) (1+v^2)}. \]

To solve the integral over \( v \) go to www.wolframalpha.com and type:
integrate \((1/(1+(u-v)^2))*(1/(1+v^2))\) from \( v=-\infty \) to \( \infty \)

The result is:
\[ \int_{-\infty}^{\infty} dv \frac{1}{(1+(u-v)^2) (1+v^2)} = \frac{2\pi}{4+u^2}. \]

This gives the CDF in integral form. Differentiating with respect to \( a \) gives the PDF:
\[ p_{X+Y}(a) = \frac{1}{\pi} \frac{2}{(4+a^2)} = \frac{1}{\pi} \frac{2}{2(1+(a/2)^2)}. \]

Thus, adding two Cauchy rv’s centered at 0 with “scale parameter” of \( \gamma = 1 \) results also in a Cauchy\textsuperscript{12} distribution centered at 0, but with \( \gamma = 2 \) (i.e. it is twice as broad as the \( \gamma = 1 \) case).

2.27. Central Limit Theorem and Applications

One of the most important theorems in probability theory is the central limit theorem (CLT). The CLT describes many important physical phenomena observed in nature that arise from the sum of many independent random effects (e.g. microscopic forces). Loosely speaking, the central limit theorem states that regardless of the distribution of these random effects, the limiting distribution is Gaussian.

Let \((X_1, \ldots, X_n)\) be a sequence of iidrv (independent identically distributed random variables), each having mean \( \mu_X \) and variance \( \sigma^2 \). Then,
\[ \lim_{n \to \infty} \text{rv} \left\{ \frac{1}{n} \sum_{i=1}^{n} (X_i - \mu_X) \right\} \xrightarrow{d} \frac{1}{\sqrt{n}} N(0, \sigma^2). \]

\textsuperscript{12}The general Cauchy distribution has PDF \( p(x) = \left[ \pi \gamma \left( 1 + \left( \frac{x-x_0}{\gamma} \right)^2 \right) \right]^{-1} \), where \( x_0 \) is the center and \( \gamma \) is the scale parameter (related to the width).
This is equivalent to saying that the arithmetic average \( \frac{1}{n} \sum_{i=1}^{n} X_i \) converges\(^{13}\) (in distribution) to a normal law with mean \( \mu_X \) and variance \( \frac{\sigma^2}{n} \):

\[
\lim_{n \to \infty} \text{rv} \left\{ \frac{1}{n} \sum_{i=1}^{n} X_i \right\} \xrightarrow{d} N \left( \mu_X, \frac{\sigma^2}{n} \right).
\]

The quantities \( \{X_i\} \) are random variables. Their sum is also a random variable; the CLT states that the sum will be Gaussian-distributed.

Note: that the arithmetic average should also have the mean \( \mu_X \) comes as no surprise. But also that the variance \( \frac{\sigma^2}{n} \) scales as \( 1/n \) should come as no surprise if you recall the definition of standard error, which states that the error in the mean scales as \( 1/\sqrt{n} \).

The central limit theorem is very important in the physical sciences because many physical measurements yield Gaussian distributions as a result of the effects of many small additive forces. For example, the Brownian motion of a particle is the result of many small collisions with solvent molecules, yielding a Gaussian distribution for the net displacement.

The CLT is illustrated in Fig. 2.12. The histogram on the left represents the probability distribution of a single rv, \( X_1 \). The histogram in the center is the distribution of the average of two such iid rv, \( \frac{1}{2} (X_1 + X_2) \). The histogram on the right is the distribution for the average of 10 iid rv, \( \frac{1}{10} (X_1 + X_2 + \cdots + X_{10}) \). As can be seen, while each rv has a uniform (flat) distribution, as the histogram on the left shows, the more random variables we average, the closer the distribution of the average approaches a normal (bell-shaped) distribution.

### 2.27.1. Random Walk in One Dimension.

The random walk, which is an application of the CLT, is important in the physical sciences. Brownian motion is a limit of the random walk.

Let \( X_i \) be the random variable which denotes the displacement at the \( i \)-th step

\[ X_i \in \{-\sigma, +\sigma\} \]

and each outcome occurs with equal probabilities, i.e.

\[ p_\sigma = \frac{1}{2} \quad \text{and} \quad p_{-\sigma} = \frac{1}{2}. \]

\(^{13}\) Convergence in distribution means that the distribution functions converge pointwise:

\[
\lim_{n \to \infty} P(X_n \leq x) = F_n(x),
\]

where \( F_n(x) \) denotes the CDF of the normal random variable with mean \( \mu \) and variance \( \sigma^2/n \) and \( X_n = \frac{1}{n} (X_1 + \cdots + X_n) \).
These displacements at different points in time are assumed to be statistically independent. After \( n \) such steps the net displacement is

\[ S_n = X_1 + \cdots + X_n. \]

Since \( X_i = \frac{1}{2}\sigma + \frac{1}{2}(-\sigma) = 0 \) we have \( S_n = 0 \). However,

\[ \text{var}X_i = \overline{X_i^2} - \overline{X_i}^2 = \sigma^2 \]

Since the \( X_i \) are statistically independent

\[ \overline{S_n^2} = \sum_{i=1}^{n} \text{var}X_i = n\sigma^2, \]

since \( \text{var}X_i = \frac{1}{4}(\sigma^2) + \frac{1}{4}(-\sigma)^2 = \sigma^2 \). The total duration of the random walk is \( t = n\Delta t \). We have

\[ \overline{S_n^2} = \left( \frac{\sigma^2}{\Delta t} \right) t = 2Dt. \]

The quantity \( D = \frac{\sigma^2}{2\Delta t} \) is called the diffusion constant (or diffusion coefficient). \( D \) has units of length square divided by time (e.g. \( \text{m}^2/\text{s} \)).

By the central limit theorem, we have that

\[ \lim_{n \to \infty} S_n = X_1 + \cdots + X_n \sim N(0, n\sigma^2) \]

\( \lim S_n \sim N(0, n\sigma^2) \) means that its PDF is

\[ p_{S_n}(x, t) = \frac{1}{\sqrt{4\pi Dt}} \exp \left[ -\frac{x^2}{4Dt} \right]. \]
Some authors call this a “conditional probability”, i.e. the probability of finding the particle at \( x, t \) given that it was at position \( x = 0 \) at time \( t = 0 \):

\[
p_{S_n}(x, t|0, 0) = \frac{1}{\sqrt{4\pi Dt}} \exp \left[ -\frac{x^2}{4Dt} \right].
\]

where \( p_{S_n}(x, t|0, 0) \) denotes the transition probability density. More generally, a transition probability density is written \( p(x, t|y, s) \) to denote the probability of finding a particle at position \( x \) at time \( t \) given that it was initially at position \( y \) at some earlier time \( s \).

The average position of the random walker after the \( n \)-th step is zero: \( \bar{S}_n = 0 \). This means that if we repeat the random walk experiments, say 10,000 times, the average position after \( n \) steps will be zero. It does not mean that the random walker automatically returns to the origin. It is merely a statement about the random walk when the walk is repeated many times.

On the other hand, the result \( \bar{S}_n^2 = 2Dt \) about the mean square displacement being proportional to \( t \) (or root mean square displacement, \( x \), being proportional to \( \sqrt{t} \)) should be contrasted to the case of ballistic motion for which \( x = vt \) (displacement proportional to \( t \)). The different powers of \( t \) reflect the fact that in diffusional motion, there is a lot of back-and-forth, leading to a shorter displacement over time.

### 2.27.2. Random Walk in Three Dimensions.

In three dimensions, the displacement is a 3-components vector \( R = (X, Y, Z) \). If the components \( X, Y, Z \) are statistically independent of each other, the joint probability density is a product:

\[
p_{XYZ}(x, y, z, t) = p_X(x, t)p_Y(y, t)p_Z(z, t),
\]

where \( p_X(x, t) = \frac{1}{\sqrt{4\pi Dt}} \exp \left[ -\frac{x^2}{4Dt} \right] \), etc. (for each component). Applying the result of the previous section for each direction (component), we obtain the joint probability density for the 3D random walk:

\[
p_{XYZ}(x, y, z, t) = \frac{1}{(4\pi Dt)^{3/2}} \exp \left( -\frac{r^2}{4Dt} \right)
\]

where \( r^2 = x^2 + y^2 + z^2 \).

We leave it to the reader to check that in 3D the mean square displacement is

\[
\mathbb{E}(r(t)^2) = \int_{\mathbb{R}^3} r^2 \exp \left( -\frac{r^2}{4Dt} \right) d^3r = 6Dt.
\]

In the general case of \( d \) dimensions, the mean square displacement is:

\[
\mathbb{E}(r(t)^2) = \mathbb{E}(x_1^2 + \cdots + x_d^2) = 2dDt.
\]
2.28. Importance Sampling

2.28.1. Law of Large Numbers (LLN). Let \( X_1, \ldots, x_n \) be a sequence of iiddv each with mean \( \mu \). Let

\[
\overline{X}_n = \frac{1}{n} \sum_{i=1}^{n} X_i = \frac{1}{n}(X_1 + \cdots + X_n).
\]

Then, for any \( \epsilon > 0 \),

\[
\lim_{n \to \infty} \mathbb{P}(|\overline{X}_n - \mu| > \epsilon) = 0.
\]

This is the weak law of large numbers (WLLN). For a proof of the WLLN, see Problem 28. \( \overline{X}_n \) is the sample mean. The strong law of large numbers is\(^{14}\)

\[
\mathbb{P}\left( \lim_{n \to \infty} |\overline{X}_n - \mu| > \epsilon \right) = 0.
\]

Both weak and strong LLN are statements about how the sample mean converges to the real mean. However, there is an important difference: the weak LLN tells us how sequences of probabilities (\( \mathbb{P} \)) converge whereas the strong LLN tells us how the sequence of rv \( \overline{X}_n \) behaves in the limit. The CLT, on the other hand, is a much stronger\(^{15}\) statement: the sample mean (arithmetic average) rv converges (in distribution) to a normal law. The central limit theorem (CLT) should not be confused with the LLN.

2.28.2. WLLN In Words. The statement \( \lim_{n \to \infty} \mathbb{P}(|\overline{X}_n - \mu| > \epsilon) = 0 \) for any \( \epsilon > 0 \) simply means that if we take any sequence of iiddrv’s \( X_1, \ldots, X_n \), their arithmetic average tends to their mathematical expectation as \( n \to \infty \). Thus, we can approximate mathematical expectations (which may be difficult to compute, especially if the distribution is unknown), using arithmetic averages formed using experimental data. The larger \( n \) is, the better the approximation of the mathematical expectation.

Since \( X_i \) are random variables, the WLLN formulation also applies to functions of random variables, \( Y_i = f(X_i) \), since the latter are also random variables. If desired, \( f \) could be almost any formula. For example, suppose that \( X_i \) is the acceleration of a particle, \( a_i \) (i.e., \( X_i \equiv a_i \)). The force is mass times acceleration: \( F(X_i) = ma = mX_i \). The WLLN allows us to compute not only the average acceleration, \( \mathbb{E}(X_i) \), but the average force, \( \mathbb{E}(F(X_i)) \), by simply renaming \( Y_i = f(X_i) \) and applying it to the sequence \( \{Y_i\} \).

\(^{14}\)The proof of the strong law requires the Borel-Cantelli lemma, which we have not covered.

\(^{15}\)This is a stronger statement because it is a statement about the entire distribution law of a rv, not just its mean and variance.
2.28.3. Mathematical Expectation With Respect To Probability Measure. The mathematical expectation of $X$ was defined as:

$$E(X) = \int_{\mathbb{R}} xp(x)dx,$$

where $p(x)$ is the PDF of $X$ and the integral is taken over the range of $X$ (here, $\mathbb{R}$). $p(x)dx$ is the probability measure. It will be useful to use the notation $E_p(X)$ to emphasize that the PDF used to calculate the mathematical expectation is $p$. This way, there is no ambiguity as to which probability measure is used. For example, if the PDF of $Y$ is $q(y)$, then we write:

$$E_q(Y) = \int yq(y)dy.$$

(Note: $y$ is a dummy integration variable; the choice of symbol is unimportant.) The WLLN states that if $Y$ is distributed according to $q(y)$, then its mathematical expectation can be approximated by the arithmetic average:

$$E_q(Y) \approx \frac{1}{n} \sum_{i=1}^{n} Y_i,$$

where the $Y_i$’s are sampled according to the distribution $q$. If instead the PDF of $Y$ had been some other function $f$, we would have written $E_f(Y)$ for $\int yf(y)dy$. The two numerical values $E_f(Y)$ and $E_q(Y)$ can, of course, be different, since $f$ and $q$ may be different functions.

As far as the WLLN is concerned, it is meant to enable us to approximate mathematical expectations of random variables (or functions of random variables) by arithmetic averages constructed using experimental data. We present several examples below to illustrate applications of the WLLN. The WLLN is best explained by working out specific examples.

2.28.4. Numerical Integration by Monte-Carlo Method. Monte-Carlo methods can be used to estimate the numerical value of integrals. For example, suppose we want to compute the integral:

$$I = \int_{a}^{b} h(x)dx,$$

which is the same as

$$I = \int_{a}^{b} \frac{h(x)(b-a)}{(b-a)}dx = \int_{a}^{b} u(x)p(x)dx = E_p(u(X))$$

where $E_p$ denotes the mathematical expectation with respect to the PDF $p(x)$ and

$$u(x) = h(x)(b-a) \text{ and } p(x) = \frac{1}{(b-a)}.$$
Thus, $p(x)$ is the PDF of a uniformly distributed random variable. This integral can be evaluated by generating random numbers $X_1, \ldots, X_n$ that are idirv and uniformly distributed on the interval $[a, b]$. By the WLLN, the following estimator converges to $I$:

$$
\hat{I} = \frac{1}{n} \sum_{i=1}^{n} u(X_i) \to \mathbb{E}_p(u(X)) = \int_{a}^{b} u(x)p(x)dx = I,
$$

This gives us a way to evaluate integrals by generating random numbers. For multi-dimensional integrals, this method has important advantages. Namely, the generation of random numbers followed by estimation of $I$ is less computationally intensive than the direct numerical integration (by quadratures) of the multi-dimensional integral.

There is no special reason to pick the uniform distribution. In fact, any distribution $p(x)$ can be used. In some cases, special choices of $p(x)$ may be advantageous. For example, sampling the domain $[a, b]$ may be a waste of time if most of the points within that interval correspond to regions where $h(x)$ is zero or nearly zero. In that case, we instead want to sample regions of the domain where $|h(x)| > 0$ is concentrated.

In other words, let

$$
I = \int_{a}^{b} h(x)dx = \int_{a}^{b} \frac{h(x)}{p(x)} p(x)dx = \mathbb{E}_p(u(X)), \quad u(x) = \frac{h(x)}{p(x)}
$$

and choose $p(x)$ such that the “peaks” of $p$ correspond to the peaks of $h$. The numerical value of $I$ can be estimated by sampling iidrv $X_1, \ldots, X_n$ according to $p(x)$ and computing the sum:

$$
I \approx \frac{1}{n} \sum_{i=1}^{n} u(X_i) = \frac{1}{n} \sum_{i=1}^{n} \frac{h(X_i)}{p(X_i)}, \quad X_i \sim p(x).
$$

2.28.5. Change of Distribution. Suppose that $X$ is a random variable with PDF $p(x)$ and we want to calculate the average of a function, $f(X)$ of $X$. Let $q(x) > 0$ be another PDF on the same probability space as $p$. Then,

$$
\mathbb{E}_p(f(X)) = \int f(x)p(x)dx = \int \frac{f(x)p(x)}{q(x)} q(x)dx = \mathbb{E}_q \left( \frac{f(X)p(X)}{q(X)} \right).
$$

Here, $\mathbb{E}_p(f(X))$ denotes the mathematical expectation of $f$ calculated using the PDF $p(x)$ for $X$, whereas $\mathbb{E}_q \left( \frac{f(X)p(X)}{q(X)} \right)$ is the expectation of $fp/q$ calculated by associating the PDF $q(x)$ to $X$ instead.

In the first case, $X_1, \ldots, X_n$ random numbers are sampled from the distribution $p(x)$ and the integral is estimated as:

$$
\mathbb{E}_p(f(X)) \approx \frac{1}{n} \sum_{i=1}^{n} f(X_i).
$$
Convergence is assured by the LLN. In the second case, \(X_1, \ldots, X_n\) are sampled from the distribution \(q(x)\) and the integral is estimated as
\[
\mathbb{E}_q \left( \frac{f(X)p(X)}{q(X)} \right) \approx \frac{1}{n} \sum_{i=1}^{n} \frac{f(X_i)p(X_i)}{q(X_i)}.
\]

The correction factor \(p/q\) is called the likelihood ratio. This method, of course, requires us to be able to compute the ratio \(p/q\) for any value of \(X\) that we may sample. The idea is to choose an importance distribution \(q\) that leads to faster convergence than the nominal distribution \(p\). We generally want to choose \(q\) such that its spikes correspond to those of \(fp\); in fact, it can be shown that \(q\) should be proportional to \(fp\). Choosing \(q\) can be done using an “educated guess” or by random sampling of the function.

2.28.6. Calculation of Probabilities. Since probabilities are expectation values of indicator functions, the WLLN can also be used to speed up the calculation of probabilities. This is especially useful for rare events. For example, suppose we want to calculate the probability of an event \(\{X \in A\}\), where the PDF for \(X\) is \(p(x)\). From experimental measurements \(X_1, \ldots, X_n\), this would normally be approximated by
\[
\mathbb{P}(X \in A) = \mathbb{E}_p(1_{X \in A}) = \int_R 1_A(x)p(x)dx \approx \frac{1}{n} \sum_{i=1}^{n} 1_A(X_i) \frac{\# \text{ draws in } A}{n}
\]
where \(1_A(x)\) is the indicator function of \(A\), i.e. it is a function that equals 1 when \(x \in A\) and 0 otherwise. However, if \(p(x)\) is such that this event rarely happens, we are going to need \(n\) very large or else the result will be zero.

On the other hand, the WLLN enables us to reweigh the integral, if we can find a better distribution \(q(x)\) that samples values that are closer to the set \(A\):
\[
\mathbb{E}_p(1_{X \in A}) = \mathbb{E}_q \left( \frac{1_A(X)p(X)}{q(X)} \right) \approx \frac{1}{n} \sum_{i=1}^{n} 1_A(X_i) \frac{p(X_i)}{q(X_i)}.
\]

If the event is rare, the ratio \(p/q\) will be small (yielding the correct probability for the rare event), whereas the summation will count several non-zero terms, giving a more accurate answer (for the same \(n\)).

A special case of \(\mathbb{P}(X \in A)\) is the estimation of the CDF, \(\mathbb{P}(X < x)\), which can be expressed as \(\mathbb{E}_p(1_{X < x})\), and which equals to \(\mathbb{E}_q \left( \frac{1_{\{X < x\}}p(X)}{q(X)} \right)\). And if \(q(x)\) is a better distribution than \(p(x)\), we can use the latter formula, together with the WLLN, to approximate the CDF by a summation.

2.28.7. Generalization to \(d\)-Dimensions. All of the above formulas are valid in \(d\)-dimensions. \(x \in \mathbb{R}\) and \(X \in \mathbb{R}\) are replaced by \(x \in \mathbb{R}^d\) and \(X \in \mathbb{R}^d\),
respectively. Likewise, \(dx\) stands for \(d^d x\), the \(d\)-dimensional volume element in the \(d\)-dimensional integral.

2.29. Review of Calculus

2.29.1. Changing Variables Under the Integral Sign. We often need to change variables under the integral sign to express the probabilistic integral in terms of known probability densities. At the same time, we should use a convenient coordinate system where the integrals can be computed.

Suppose that we want to integrate \(f(u, v)\) over a region \(R\). Under the inverse of the transformation \(u = u(x, y), v = v(x, y)\) the region \(R\) becomes \(S\) and the double integral becomes

\[
\int\int_R f(u, v) \, du \, dv = \int\int_S f(u(x, y), v(x, y)) \left| \frac{\partial (u, v)}{\partial (x, y)} \right| \, dx \, dy,
\]

where \(\frac{\partial (u, v)}{\partial (x, y)}\) is the Jacobian determinant:

\[
\frac{\partial (u, v)}{\partial (x, y)} \equiv \begin{vmatrix} \frac{\partial u}{\partial x} & \frac{\partial u}{\partial y} \\ \frac{\partial v}{\partial x} & \frac{\partial v}{\partial y} \end{vmatrix} = \frac{\partial u}{\partial x} \frac{\partial v}{\partial y} - \frac{\partial u}{\partial y} \frac{\partial v}{\partial x}.
\]

If we look at the area elements, we see that the Jacobian plays the role of a scaling factor indicating by how much the area element is scaled in the new coordinates:

\[
du \, dv = \left| \frac{\partial (u, v)}{\partial (x, y)} \right| \, dx \, dy.
\]

This idea extends to multi-dimensional integrals, where the Jacobian represents the scaling of the volume element, etc. To avoid cluttering, I will often use the shorthand notation \(\frac{\partial u}{\partial x} = \frac{\partial}{\partial x}\) etc. for partial derivatives.

For those who don’t remember the change-of-variables formula, there is another method which does not require you to remember anything except for the alternating algebra of differential forms. \(du\) and \(dv\) are differential 1-forms (covector fields). The product \(du \wedge dv\) should instead be viewed as a “wedge product” \(du \wedge dv\), which is a 2-form. A 2-form \(du \wedge dv\) is an oriented area element spanned by the covectors \(du\) and \(dv\). In the left hand side we replace \(du \wedge dv\) by the wedge product \(du \wedge dv\)

\[
\int\int_R f(u, v) \, du \, dv = \int\int_R f(u, v) \, du \wedge dv
\]

Then, viewing \(u\) and \(v\) as functions of \(x\) and \(y\), we expand \(du\) and \(dv\) as total differentials:

\[
du(x, y) = (\partial_x u) dx + (\partial_y u) dy
\]

and

\[
dv(x, y) = (\partial_x v) dx + (\partial_y v) dy.
\]
Then we form the wedge product of $du$ and $dv$:

$$du \wedge dv = [(\partial_x u)dx + (\partial_y u)dy] \wedge [(\partial_x v)dx + (\partial_y v)dy],$$

When distributing the product, we apply the rules of the alternating algebra: since $dx \wedge dy$ is an oriented area element spanned by the covectors $dx$ and $dy$, we have that $dx \wedge dy = -dy \wedge dx$ (sign flipped because we have an “oriented area element” and this amounts to changing from the left hand rule to the right hand rule in a cross product), $dx \wedge dx = 0$ and $dy \wedge dy = 0$ (zero because the area spanned by two collinear vectors is zero). We are left with:

$$du \wedge dv = (\partial_x u)(\partial_y v)dx \wedge dy + (\partial_x u)(\partial_y v)dy \wedge dx = (\partial_x u \partial_y v - \partial_y u \partial_x v)dx \wedge dy.$$

You will recognize the coefficient of $dx \wedge dy$ on the right hand side as the Jacobian determinant $\frac{\partial(u,v)}{\partial(x,y)}$. Thus, the alternating algebra of differential forms took care of calculating the determinant for us. This works in any number of dimensions.

Let us work out an example. Suppose that we have an integral

$$\int \int_R f(x,y) \, dx \, dy$$

and want to change from Cartesian to polar coordinates, i.e.

$$x = r \cos \theta, \quad y = r \sin \theta.$$

The total differentials are:

$$dx(r, \theta) = \partial_r x \, dr + \partial_\theta x \, d\theta = \cos \theta \, dr - \sin \theta \, d\theta$$

$$dy(r, \theta) = \partial_r y \, dr + \partial_\theta y \, d\theta = \sin \theta \, dr + \cos \theta \, d\theta.$$

Forming the wedge product $dx \wedge dy$:

$$dx \wedge dy = [\cos \theta \, dr - \sin \theta \, d\theta] \wedge [\sin \theta \, dr + \cos \theta \, d\theta].$$

Applying the rules $d\theta \wedge d\theta = 0$, $dr \wedge dr = 0$ and $dr \wedge d\theta = -d\theta \wedge dr$, we are left with:

$$dx \wedge dy = r \cos^2 \theta \, dr \wedge d\theta - r \sin^2 \theta \, d\theta \wedge dr = r \, dr \wedge d\theta,$$

which is the familiar area element in polar coordinates.

Now let us return to the example of the previous section where we had the integral $\int \int p_U(u)p_V(v)\, du\, dv$, where $v = z$ and $u = zy$. Writing $du\, dv$ as a wedge product $du \wedge dv$, expanding the total differentials: $du(y,z) = zdy+ydz$ and $dv(y,z) = dz$ yields $du\wedge dv = zdy\wedge dz$, where $z$ is the Jacobian determinant that was sought and $dy \wedge dz$ are the new integration variables. The alternating algebra of differential forms automatically computes the determinant for us.
2. Probability

2.29.2. Leibniz Formula. The Leibniz formula for differentiation of integrals (the Leibniz integral rule) is:
\[
\frac{d}{dy} \left( \int_{a(y)}^{b(y)} f(x, y) \, dx \right) = \int_{a(y)}^{b(y)} \frac{\partial}{\partial y} f(x, y) \, dx + f(b(y), y) \cdot b'(y) - f(a(y), y) \cdot a'(y)
\]
which consists of the sum of three terms: in the first one the differentiation is carried out inside the integral; the remaining two terms are surface (boundary) terms. This formula will help you compute PDFs from CDFs.

Let’s look at an example of obtaining the PDF from the CDF \( P(Y < y) \) when \( X \sim N(\mu, \sigma^2) \) and \( Y = e^X \), by differentiating with respect to \( y \). Since the \( y \) dependence occurs only in the upper limit of the integral, only the second term in the Leibniz formula is non-zero:
\[
p_Y(y) \equiv \frac{d}{dy} P(Y < y) = \frac{d}{dy} \int_{-\infty}^{\log y} \frac{1}{\sqrt{2\pi\sigma^2}} e^{-(x-\mu)^2/2\sigma^2} \, dx
\]
\[
= \frac{1}{\sqrt{2\pi\sigma^2}} e^{-(\log y-\mu)^2/2\sigma^2} \cdot \frac{1}{y}.
\]
This particular PDF is also known as the log-normal distribution.

2.29.3. Infinitesimals. It is important to know how to work with infinitesimals. Let us take the link between CDF and PDF as an example. Denoting the CDF as \( \mathbb{P}(X < x) = F(x) \), let us Taylor expand \( F(x + \epsilon) \) about the point \( x \) (here \( \epsilon > 0 \) is a small quantity):
\[
(2.12) \quad F(x + \epsilon) = F(x) + \epsilon \cdot F'(x) + o(|\epsilon|)
\]
where \( F'(x) = dF(x)/dx \) and \( o(|\epsilon|) \) denotes higher order terms (in this case \( \epsilon^2 \) and higher powers of \( \epsilon \)) which decay to 0 faster than \( \epsilon \) in the limit \( \epsilon \to 0 \):
\[
\lim_{\epsilon \to 0} \frac{o(|\epsilon|)}{|\epsilon|} = 0,
\]
so that taking the limit \( \epsilon \to 0 \) in Eq. (2.12) leads to \( F(x + \epsilon) = F(x) + \epsilon dF(x)/dx \). Taking \( \epsilon = dx \) (infinitesimal) this can be rewritten as:
\[
F(x + dx) - F(x) = F'(x) \cdot dx + o(|dx|) = F'(x) \cdot dx + o(|dx|),
\]
since \( F'(x) = dF(x)/dx \). Thus, as \( dx \to 0 \) (without being equal to 0) the term \( o(|dx|) \) vanishes and we have that
\[
(2.13) \quad F(x + dx) - F(x) = dF(x).
\]
Denote the CDF as \( F(x) \equiv \mathbb{P}(X < x) \) and recall the interpretation of the PDF. Given a random variable \( X \) its PDF \( p_X(x) \) times \( dx \) gives the probability that \( X \) will lie in the interval \( (x, x+dx) \):

\[
p_X(x)dx = \mathbb{P}(x \leq X \leq x + dx) = \mathbb{P}(X \leq x + dx) - \mathbb{P}(X \leq x) = d\mathbb{P}(X \leq x),
\]

where \( dF(x) = F(x+dx) - F(x) \) was used in the last step. In the second equality we have made use of the definition of the “probability function” \( \mathbb{P}(\cdot) \) as an integral of the PDF, i.e.

\[
\mathbb{P}(a \leq X \leq b) = \int_a^b p_X(x)dx = \int_{-\infty}^b p_X(x)dx - \int_{-\infty}^a p_X(x)dx = \mathbb{P}(X \leq b) - \mathbb{P}(X \leq a)
\]

hence \( \mathbb{P}(x \leq X \leq x + dx) = \mathbb{P}(X \leq x + dx) - \mathbb{P}(X \leq x) \). Another way to look at the quantity \( \mathbb{P}(x \leq X \leq x + dx) \) is the integral of the PDF from \( x \) to \( x + dx \)

\[
\mathbb{P}(x \leq X \leq x + dx) = \int_x^{x+dx} p_X(x')dx' = p_X(x)dx.
\]

The last equality follows because the integral is a Riemann sum containing only 1 term. It contains only 1 term because the interval \( [x, x+dx] \) where the integral is carried out is infinitesimally small.

So integrating from \( a \) to \( b \) we get the probability that \( X \) takes values between \( a \) and \( b \):

\[
\int_a^b p_X(x')dx' = \int_a^b \mathbb{P}(x \leq X \leq x + dx) = \int_a^b [\mathbb{P}(X \leq x + dx) - \mathbb{P}(X \leq x)]
\]

\[
= \int_a^b d\mathbb{P}(X \leq x) = \mathbb{P}(X \leq b) - \mathbb{P}(X \leq a)
\]

\[
= \mathbb{P}(a \leq X \leq b).
\]

If the interval \( (a, b) \) is chosen to be \( (-\infty, x) \) we get the CDF:\(^{16}\)

\[
\int_{-\infty}^x p_X(x')dx' = \mathbb{P}(-\infty \leq X \leq x) = \mathbb{P}(X \leq x).
\]

Differentiating with respect to \( x \) yields a method for obtaining the PDF from the CDF:

\[
\frac{d\mathbb{P}(X \leq x)}{dx} = p_X(x).
\]

\(^{16}\)Notice that we wrote \( \mathbb{P}(X \leq x) \) instead of \( \mathbb{P}(-\infty \leq X \leq x) \) because the statement that \( X \geq -\infty \) is always true and therefore, unnecessary or redundant.
So now you know how to go from PDF to CDF or from CDF to PDF. The two concepts are related to each other by an integral or a derivative. If you are asked to obtain the probability distribution of a rv you can derive either the PDF or the CDF. In general, obtaining the CDF is easier because fewer steps are needed and the interpretation of the CDF in terms of probability is also simpler.

2.30. Problems

Problem 20. The Poisson’s law with parameter $a$ ($a > 0$) is defined by:

$$
P[k \text{ events}] = e^{-a} \frac{a^k}{k!},
$$

where $k = 0, 1, 2, \ldots$. With $a = \lambda \tau$, where $\lambda$ is the average number of events per units time and $\tau$ is the length of the interval $(t, t + \tau)$, the probability of $k$ events in $\tau$ is

$$
P(k; t, t + \tau) = e^{-\lambda \tau} \frac{(\lambda \tau)^k}{k!}.
$$

This equation assumes that $\lambda$ is independent of $t$. If $\lambda$ depends on $t$, the product $\lambda \tau$ gets replaced by the integral $\int_t^{t+\tau} \lambda(\xi) d\xi$, and the probability of $k$ events in the interval $(t, t + \tau)$ is

$$
P(k; t, t + \tau) = \exp \left[ - \int_t^{t+\tau} \lambda(\xi) d\xi \right] \frac{1}{k!} \left[ \int_t^{t+\tau} \lambda(\xi) d\xi \right]^k.
$$

The parameter $\lambda$ is called the rate parameter. $\lambda(t)$ is the rate function. Suppose that a company manufactures superconducting wire. Upon close examination of the product on the assembly line, you find that the defect density along the length of the wire is not uniform. For wire strips of length $D$, the defect density $\lambda(x)$ along the wire length $x$ varies as

$$
\lambda(x) = \lambda_0 + \frac{1}{2}(\lambda_1 - \lambda_0) \left( 1 + \cos\left( \frac{2\pi x}{D} \right) \right), \quad \lambda_1 > \lambda_0
$$

for $0 \leq x \leq D$ due to greater wire contamination at the edges $x = 0$ and $x = D$.

(i) What is the meaning of $\lambda(x)$ in this case?

(ii) What is the average number of defects for a wire strip of length $D$?

(iii) Find an expression for the probability of $k$ defects on a wire strip of length $D$?

Solution. (i) Bearing in mind that $\lambda(x)$ is a defect density, i.e., the average number of defects per unit length at $x$, we conclude that $\lambda(x) \Delta x$ is the average number of defects in the tape from $x$ to $x + \Delta x$. 
2.30. Problems

(ii) Given the definition of $\lambda(x)$ we conclude that the average number of defects along the whole wire is merely the integral of $\lambda(x)$, i.e.,

$$\int_0^D \lambda(x)dx = \int_0^D \left[ \lambda_0 + \frac{1}{2}(\lambda_1 - \lambda_0) \left( 1 + \cos \frac{2\pi x}{D} \right) \right] dx = \frac{\lambda_0 + \lambda_1}{2} D = \Omega.
$$

(iii) Assuming the Poisson law holds, use the equation with $x$ and $\Delta x$ (distances) replacing $t$ and $\tau$ (times). Thus,

$$P(k; x, x+\Delta x) = \exp\left(-\int_x^{x+\Delta x} \lambda(\zeta)d\zeta\right) \frac{1}{k!} \left(\int_x^{x+\Delta x} \lambda(\zeta)d\zeta\right)^k.
$$

In particular, with $x = 0$ and $x + \Delta x = D$, we obtain

$$P(k; 0, D) = \Omega^k \frac{e^{-\Omega}}{k!}.
$$

Problem 21. Let $X$ and $Y$ be independent random variables having the exponential distribution with parameters $\lambda$ and $\mu$ respectively. (Recall that if rv $X$ has exponential distribution with parameter $\lambda > 0$, its CDF is $P(X < x) \equiv 1 - e^{-\lambda x}$, $x \geq 0$, whose density is $dP(X < x)/dx = \lambda e^{-\lambda x}$.) Let $U = \min\{X, Y\}$, $V = \max\{X, Y\}$ and $W = V - U$. Find the probability $P(U = X) = P(X \leq Y)$. Show that $U$ and $W$ are statistically independent.

Solution. First you should realize that the logical statements $U = X$ and $X \leq Y$ mean the same thing. Indeed, $U = \min\{X, Y\}$ less than or equal to $X$ implies that both $X \leq X$ (if $\min\{X, Y\} = X$) and $Y \leq X$ (if $\min\{X, Y\} = Y$). The former ($X \leq X$) is a trivial statement which is true at all times. Thus, it can be ignored. The only non-trivial statement left is $Y \leq X$, hence the equivalence of the two statements $U = X$ and $X \leq Y$. If the two statements are equivalent, then their probabilities are also equal: $P(U = X) = P(X \leq Y)$. 

\begin{figure}[h]
\centering
\includegraphics[width=0.5\textwidth]{figure2.13.png}
\caption{Integration in the $x < y$ region.}
\end{figure}
\[
\mathbb{P}(U = X) = \mathbb{P}(X \leq Y) \text{ can be computed since it is in terms of } X \text{ and } Y \text{ whose distributions are known. Since } \mathbb{P}(X \leq Y) \text{ involves both } X \text{ and } Y \text{ we must integrate the joint PDF of } X \text{ and } Y \text{ over the set of all points } (x, y) \text{ such that } x < y \text{ is satisfied:}
\]

\[
\mathbb{P}(X \leq Y) = \int_{\{(x,y) | x < y\}} p_{XY}(x,y) \, dx \, dy.
\]

Let’s integrate along horizontal strips, as shown in Fig. 2.13. Thus,

\[
\mathbb{P}(X \leq Y) = \int_0^\infty \mathbb{P}(X \leq y) \mu e^{-\lambda y} 
\]

\[
\int_0^\infty (1 - e^{-\lambda y})\mu e^{-\mu y} 
\]

\[
\int_0^\infty \lambda e^{-\lambda x} e^{-\mu(x+w)} 
\]

\[
\lambda e^{-\mu} (1 - e^{-(\lambda+\mu)u})
\]

and similarly, \( \mathbb{P}(U \leq u, W > w, X > Y) = \frac{\mu}{\lambda + \mu} e^{-\lambda w} (1 - e^{-(\lambda+\mu)u}) \). Hence, for \( 0 \leq u \leq u + w < \infty \), we have an expression which factorizes into the product of a function of \( u \) with a function of \( w \). Hence \( U \) and \( W \) are independent:

\[
\mathbb{P}(U \leq u, W > w) = (1 - e^{-(\lambda+\mu)u}) \left( \frac{\lambda}{\lambda + \mu} e^{-\mu w} + \frac{\mu}{\lambda + \mu} e^{-\lambda w} \right).
\]

\[\text{Since the two events } \{U \leq u, W > w, X \leq Y\} \text{ and } \{U \leq u, W > w, X > Y\} \text{ are mutually exclusive whereas the event } \{X \leq Y\} \cup \{X > Y\} \text{ is always true. Recall that two events } A \text{ and } B \text{ are mutually exclusive if there is no overlap: } A \cap B = \emptyset.\]
Problem 22. A coin is flipped \(n\) times. The outcome is a random variable \(X\), which can take the value heads or tails (\(X = \text{heads}\) or \(X = \text{tails}\)). For \(n\) measurements, there are \(n\) such random variables (and corresponding results): \(X_1, X_2, \ldots, X_n\). The coin is possibly biased. Therefore, the probabilities of each outcome are not necessarily 1/2. Instead they are given in term of a parameter \(-1/2 \leq \theta \leq 1/2\) quantifying the bias:
\[
\mathbb{P}(X = \text{heads}) = 1/2 + \theta, \quad \mathbb{P}(X = \text{tails}) = 1/2 - \theta.
\]
(a) Explain how the numerical value of the bias parameter, \(\theta\), can be determined experimentally (empirically) by flipping the coin several times, i.e. find an explicit formula for \(\hat{\theta}_n\), the estimator of \(\theta\), in terms of \(X_1, \ldots, X_n\). Show that, under specific circumstances, \(\hat{\theta}_n\) converges to \(\theta\) in probability, i.e. show that \(\mathbb{P}(|\hat{\theta}_n - \theta| > \epsilon) \to 0\), as \(n \to \infty\) for any \(\epsilon > 0\).
(b) Denote the number of times where you get heads as the result by \(H\) (and \(H\) is a random variable, because its value may differ each time this experiment is done). Write down an explicit expression for \(H\) in terms of the experimental data. Find the mathematical expectation of \(H\).
(c) Find the variance of \(H\). For which value(s) of \(\theta\) is the variance a minimum? A maximum?
(d) Calculate the “signal-to-noise ratio” of \(H\). Explicitly give the dependence of SNR on \(n\).
(e) For a fixed value of \(n\), find the conditions for which the SNR is 1) infinite and 2) undetermined/undefined. Give a physical explanation of those two different situations.
(f) Find the limiting (\(n\) large) distribution of \(H\).

Solution. (a) The probability \(\mathbb{P}(X = \text{heads})\) can be determined by counting the number of heads, i.e. let \(f_H\) be the empirical probability
\[
f_H = \frac{1}{n} \sum_{i=1}^{n} 1_{\{X_i = \text{heads}\}},
\]
where \(1_{\{X_i = \text{heads}\}}\) equals 1 if \(X_i = \text{heads}\) and 0 otherwise. Taking the mathematical expectation we get
\[
\mathbb{E}f_H = \frac{1}{n} \sum_{i=1}^{n} \mathbb{E}1_{\{X_i = \text{heads}\}},
\]
where
\[
\mathbb{E}1_{\{X_i = \text{heads}\}} = \sum_{\{x_i : x_i = \text{heads}\}} \mathbb{P}(X_i = \text{heads}) = \mathbb{P}(X_i = \text{heads})
\]
Therefore (the \( X_i \) are iidrv, with the same distribution as \( X \)),
\[
\mathbb{E} f_H = \mathbb{P}(X = \text{heads}).
\]

By the law of large numbers (LLN), \( f_H \) converges to \( \mathbb{P}(X = \text{heads}) \) as \( n \) increases. Now, since \( \mathbb{P}(X = \text{heads}) = 1/2 + \theta \), which is also equal to \( \mathbb{E} f_H \), we take our estimator \( \hat{\theta} \) to be:
\[
\hat{\theta}_n = \left( \frac{1}{n} \sum_{i=1}^{n} \mathbf{1}_{\{X_i = \text{heads}\}} \right) - \frac{1}{2},
\]
which implies that \( \hat{\theta}_n \) converges to \( \mathbb{P}(X = \text{heads}) - \frac{1}{2} \), as \( n \) increases. However, \( \mathbb{P}(X = \text{heads}) - \frac{1}{2} \) is also equal to \( \theta \), by the LLN. Thus, \( \hat{\theta}_n \to \theta \).

(b) \[
H = \sum_{i=1}^{n} \mathbf{1}_{\{X_i = \text{heads}\}}
\]
where \( \mathbf{1}_{X_i = \text{heads}} \) equals 1 when \( X_i = \text{heads} \) and 0 when \( X_i = \text{tails} \). Taking expectation value:
\[
\mathbb{E} H = \sum_{i=1}^{n} \mathbb{E} \mathbf{1}_{\{X_i = \text{heads}\}} = \sum_{i=1}^{n} \mathbb{P} \{ \{X_i = \text{heads}\} \} = \sum_{i=1}^{n} (1/2 + \theta) = n(1/2 + \theta).
\]

(c) Variance:
\[
\text{var}(H) = \sum_{i=1}^{n} \text{var}(\mathbf{1}_{\{X_i = \text{heads}\}}) = \sum_{i=1}^{n} \mathbb{E} [(\mathbf{1}_{\{X_i = \text{heads}\}})^2] - [\mathbb{E} (\mathbf{1}_{\{X_i = \text{heads}\}})]^2
\]
\[
= n(1/2 + \theta) - n(1/2 + \theta)^2.
\]

since \( (\mathbf{1}_{\{X_i = \text{heads}\}})^2 = \mathbf{1}_{\{X_i = \text{heads}\}} \). The variance reaches a maximum when \( \theta = 0 \) and a minimum when \( \theta = \pm 1/2 \).

(d) Find the dependence of SNR on \( n \):
\[
\text{SNR} = \frac{n(1/2 + \theta)}{\sqrt{n} \sqrt{1/4 - \theta^2}} \propto \sqrt{n}.
\]

(e) SNR is undetermined when \( \theta = -1/2 \) (probability of heads=0). 1) SNR is infinite when \( \theta = 1/2 \) (probability of heads=1).

(f) By the CLT, the limiting distribution is Gaussian. The mean is \( n(1/2 + \theta) \) and variance is \( n(1/2 + \theta) - n(1/2 + \theta)^2 \).

Problem 23. Consider a die which is unbiased. (a) You roll the die once. What is the probability of getting a “1” vs a “4”? 

(b) You roll the die twice. What is the probability of getting a total of “2” (i.e. “1” on both trials) versus the probability of getting a total of “7” (“Total” means you add the two results together.)
(c) You roll the die 10,000 times and record the results. What is the probability distribution of the mean (i.e. the mean of all the results), its first moment and variance?

**Solution.** (a) 1/6 and 1/6
(b) 2: 1/6 times 1/6 = 1/36
7: 6 times 1/6 times 1/6 = 1/6
(c) by the CLT the distribution converges to the normal law, $N(3.5, \sigma^2/10000)$, where the value of $\sigma^2$ is:

$$\sigma^2 = \sum_{i=1}^{6} (x_i - \mu)^2 p_i = 2.917$$

\[\square\]

**Problem 24.** What power of $t$ (time) does the root-mean-square displacement in a 1D random walk depend on? How does this differ from the case of ballistic motion. Explain.

**Solution.** For random walk the root mean square displacement is proportional to $\sqrt{t}$ whereas for ballistic motion it depends on $t$. The $\sqrt{t}$ dependence can be explained by the large number of “back-and-forth” steps in the random walk.

\[\square\]

**Problem 25.** Consider the normal (Gaussian) distribution with parameters $\mu$ and $\sigma^2$, i.e. let $X \sim N(\mu, \sigma^2)$. Show all calculations.

(a) Calculate moments of all orders ($n = 0, 1, 2, 3, \ldots$) for $X$.
(b) Calculate the central moments of all orders for $X$.
(c) Define a new function $K(t) = \log \mathbb{E}(e^{tX})$, and Taylor expand $K(t)$ in powers of $t$:

$$K(t) = \sum_{n=1}^{\infty} \kappa_n \frac{t^n}{n!}.$$ Find a general expression for the coefficients $\kappa_n$.

(d) Define a new function $M(t) = \exp(K(t))$. Show how the moments can be obtained from $M(t)$ in terms of the $\kappa_n$'s.

(e) Show how the density of $X$, $p(x)$, can be constructed from a knowledge of the statistical moment, or from the central moments, or from the $\kappa_n$'s.

(f) Explain why the method in (e) of reconstructing $p(x)$ is important from an experimental science standpoint.

**Solution.** (a) The moments of odd orders are all zero because the integral of an odd function ($n$-th moment of $X$, where $n$ is odd) times an even function (Gaussian PDF) vanishes because the integrand is odd. On the other hand,
the moment $\mathbb{E}(X^n)$, where $n$ is even are non-zero. They are calculated as follows:

$$\mathbb{E}(X^n) = \int_{\mathbb{R}} \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{(x-\mu)^2}{2\sigma^2}} \cdot x^n \, dx,$$

where we use

$$\int_{0}^{\infty} e^{-ax^2} x^n \, dx = \frac{(n-1)!!}{2^{n/2+1/2} a^{n/2}} \sqrt{\frac{\pi}{a}}$$

for $n$ even. The result is:

$$\mathbb{E}(X^n) = \sigma^n (n-1)!!.$$

The first few moments are:

<table>
<thead>
<tr>
<th>order</th>
<th>moment</th>
<th>central moment</th>
</tr>
</thead>
<tbody>
<tr>
<td>$n = 1$</td>
<td>$\mu$</td>
<td>0</td>
</tr>
<tr>
<td>$n = 2$</td>
<td>$\mu^2 + \sigma^2$</td>
<td>$\sigma^2$</td>
</tr>
<tr>
<td>$n = 3$</td>
<td>$\mu^3 + 3\mu\sigma^2$</td>
<td>0</td>
</tr>
<tr>
<td>$n = 4$</td>
<td>$\mu^4 + 6\mu^2\sigma^2 + 3\sigma^4$</td>
<td>$3\sigma^4$</td>
</tr>
<tr>
<td>$n = 5$</td>
<td>$\mu^5 + 10\mu^3\sigma^2 + 15\mu\sigma^4$</td>
<td>0</td>
</tr>
</tbody>
</table>

see [https://en.wikipedia.org/wiki/Normal_distribution#Moments](https://en.wikipedia.org/wiki/Normal_distribution#Moments)

(b) See solution to (a).

(c) Expand $e^{tX} = \sum_{i=0}^{\infty} \frac{t^i X^i}{i!}$ and take the average, $\mathbb{E}e^{tX} = \sum_{i=0}^{\infty} \frac{t^i \mathbb{E}(X^i)}{i!}$.

On the other hand, take the exponential of $K(t)$,

$$e^{K(t)} = e^{\sum_{n=1}^{\infty} \kappa_n \frac{t^n}{n!}} = 1 + \sum_{n=1}^{\infty} \kappa_n \frac{t^n}{n!} + \frac{1}{2} \left( \sum_{n=1}^{\infty} \kappa_n \frac{t^n}{n!} \right)^2 + \frac{1}{3!} \left( \sum_{n=1}^{\infty} \kappa_n \frac{t^n}{n!} \right)^3 + \ldots .$$

We can now identify the like powers of $t$ (let $\mu_r = \mathbb{E}X^r$):

- $t^1: \mu_1 = \kappa_1$, $\mu_1 = \kappa_1$
- $t^2: \frac{\mu_2}{2} = \frac{\kappa_2}{2} + \frac{\kappa_1^2}{2}$, $\mu_2 = \kappa_2 + \frac{\kappa_1^2}{2}$
- $t^3: \frac{\mu_3}{6} = \frac{\kappa_3}{6} + \kappa_1 \kappa_2 / 2 + \frac{\kappa_1^3}{6}$, $\mu_3 = \kappa_3 + 3 \kappa_1 \kappa_2 + \frac{\kappa_1^3}{6}$
- $\vdots$

This can be “inverted” to give:

- $\kappa_1 = \mu_1$
- $\kappa_2 = \mu_2 - \frac{\mu_1^2}{2}$
- $\kappa_3 = \mu_3 - 3\mu_2 \mu_1 + 2 \mu_1^3$
- $\vdots$
(d) \( M(t) = \mathbb{E}(e^{tX}) = \mathbb{E} \sum_{i=0}^{\infty} t^i X^i / i! = \sum_{i=0}^{\infty} t^i \mathbb{E}(X^i) / i! \). The moments are obtained by differentiation with respect to \( t \) and setting \( t = 0 \):

\[
\mu_r = \mathbb{E}(X^r) = \left. \frac{d^r}{dt^r} M(t) \right|_{t=0}.
\]

(e) Consider the quantity \( \mathbb{E}(e^{tX}) \),

\[
M(t) = \sum_{r=0}^{\infty} \frac{t^r \mu_r}{r!} = \mathbb{E}(e^{tX}) \equiv \int_{-\infty}^{\infty} e^{tx} p(x) dx.
\]

We can solve for \( p(x) \) by invoking the inversion formula:

\[
\frac{1}{2\pi i} \lim_{T \to \infty} \int_{\gamma - iT}^{\gamma + iT} e^{-st} M(s) ds = \frac{1}{2\pi i} \lim_{T \to \infty} \int_{\gamma - iT}^{\gamma + iT} e^{-st} \left[ \int_{-\infty}^{\infty} e^{sx} p(x) dx \right] ds.
\]

Then integrating over \( \tau \) and invoking \( \delta(x - a) = \frac{1}{2\pi} \int_{-\infty}^{\infty} e^{i(x-a)t} dt \):

\[
\int_{-\infty}^{\infty} e^{\gamma(x-t)} \delta(x-t)p(x) dx = p(t).
\]

(f) This is important because if we know all the moments, we can reconstruct \( p(x) \). The moments can be estimated from experimental data.

**Problem 26.** The scores in a chemistry class from 2016 were as follows (out of 100):


(a) If the course policy is to assign 'A' grades to the top 10-percentile of the class and 'F' to the rest, how many students obtained an F?

(b) Draw a histogram of the exam results.

(c) Calculate the mean, standard deviation and median of the exam and indicate those quantities on the histogram. (Explain how those quantities are calculated from the data.)

(d) In units of standard deviation (\( \sigma \)), how far is the 10-percentile from the mean?

(e) Reconstruct the PDF of this random variable \( X \) (score), using the numerical data.
(f) Suppose that the 2017 scores were:

59.378 102.006 54.660 39.713 46.731 17.570 45.646 71.654 19.959 
58.948 57.506 78.838 31.859 20.175 31.766 39.408 41.096 31.092 52.754 53.712 
67.778 66.991 37.362 57.768 72.032 48.005 78.559 46.742 84.157 66.175 90.976 
72.627 40.335 19.464 60.673 51.911 34.235 35.143 39.269 48.814 83.537 50.505 
40.340 47.480 58.682 72.354 56.195 74.103 50.013

Plot histogram and calculate distribution parameters. Are these scores significantly different from those of 2016? (Why/why not?)

Solution. (a) The top 10-percentile is called the 90th percentile, and is the value of $x_{90}$ such that

$$\int_{-\infty}^{x_{90}} p(x)dx = 0.90.$$ 

Numerically, there is a function in MATLAB called `prctile` that will compute this for us. I get 55.68 for the above data. This means that all scores below this get an 'F'; we count 45 of those.

(b) See Figure 2.14.

(c) Mean = 41.679 (use formula for sample mean), std = 10.372 (use formula for sample standard deviation), median = 40.177 (order the numbers and pick the middle one).

(d) In MATLAB, we simply type `(prctile(d,90)-mean(d))/std(d)` and obtain 1.3500.
(e) From the raw data we can calculate the moments of the distribution:

<table>
<thead>
<tr>
<th>$r$</th>
<th>raw moment</th>
<th>central moment</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>41.679</td>
<td>9.4502 × 10^{-15}</td>
</tr>
<tr>
<td>2</td>
<td>1842.5</td>
<td>105.43</td>
</tr>
<tr>
<td>3</td>
<td>8.5908 × 10^4</td>
<td>325.47</td>
</tr>
<tr>
<td>4</td>
<td>4.1974 × 10^6</td>
<td>2.6805 × 10^4</td>
</tr>
<tr>
<td>5</td>
<td>2.1355 × 10^8</td>
<td>2.0898 × 10^5</td>
</tr>
</tbody>
</table>

and using the inversion formula, we can obtain $p(x)$.

(f) The mean/std are $53.371 \pm 19.136$. Compare this to $41.679 \pm 10.372$. These two numbers are not significantly different because their error bars overlap considerably.

Problem 27. Prove, using the law of large numbers, that the histogram of a random variable $X$ converges to its PDF, $p(x)$.

Solution. Let $X$ have CDF $F(x)$. Let $X_1, X_2, \ldots, X_n$ be a random sample of $F$. Define the indicator function $\mathbb{I}_{(-\infty,x]}(y)$ to be equal to 1 if $y \leq x$ and zero otherwise. Then,

$$
\mathbb{E} \mathbb{I}_{(-\infty,x]}(X_i) = \int_{-\infty}^{x} p(x_i) dx_i = \mathbb{P}(X_i \leq x) = F(x).
$$

For each $n$, the histogram of the random sample is:

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

Its expectation value is:

$$
\mathbb{E} F_n(x) = \frac{1}{n} \sum_{i=1}^{n} \mathbb{E} \mathbb{I}_{(-\infty,x]}(X_i) = F(x).
$$

It then follows from the law of large numbers that $F_n(x)$ converges to $F(x)$. If the CDFs converge, the PDFs also converge.

Problem 28. Consider the weak law of large numbers (WLLN): Let $X_1, X_2, \ldots$ be iidrv with mean $\mu$ and variance $\sigma^2 < \infty$. Then, $(1/n) \sum_{i=1}^{n} X_i$ converges to $\mu$ in probability.

(a) Prove the WLLN.

(b) Illustrate it using a numerical example, i.e. choose $\epsilon > 0$, generate random sequences $X_1, \ldots, X_n$, compute the sample mean $\bar{X}_n$, record this value as $m_1$. Generate a second random sequence, and obtain the sample mean as $m_2$. Repeat this process many times ($r$ times) and plot a histogram of the sample means $(m_1, m_2, \ldots, m_r)$. Then increase $n$ and repeat this process.
process. You should now have several histograms as function of \( n \). Then plot the probability \( \mathbb{P}(|\bar{X}_n - \mu| \geq \epsilon) \) as a function of \( n \) and show that it converges to 0 as \( n \) increases. Since we are dealing with experimental data, the probability should be calculated empirically:

\[
\mathbb{P}(|\bar{X}_n - \mu| \geq \epsilon) = \frac{1}{r} \sum_{j=1}^{r} I_{|m_j - \mu| \geq \epsilon}
\]

where \( I_{|m_j - \mu| \geq \epsilon} \) is an “indicator function”, i.e. equals 1 when \(|m_j - \mu| \geq \epsilon\) and equals 0 otherwise.

**Solution.** (a) Weak law: let \( \bar{X}_n = (1/n)(X_1 + X_2 + \ldots + X_n) \), \( \text{var}(\bar{X}_n) = (1/n^2) \cdot \text{var}(X_1) = \sigma^2/n \), and \( \bar{x}_n = \frac{x_1 + \ldots + x_n}{n} \). Then,

\[
\mathbb{P}(|\bar{X}_n - \mu| \geq \epsilon) = \int_{\{|x_1,\ldots,x_n|:|\bar{x}_n - \mu| \geq \epsilon\}} p_1(x_1) \ldots p_n(x_n) dx_1 \ldots dx_n
\]

\[
\leq \int_{\{|\bar{x}_n - \mu| \geq \epsilon\}} \frac{(\bar{x}_n - \mu)^2}{\epsilon^2} p_1(x_1) \ldots p_n(x_n) dx_1 \ldots dx_n
\]

\[
\leq \int_{\mathbb{R}^n} \frac{(\bar{x}_n - \mu)^2}{\epsilon^2} p_1(x_1) \ldots p_n(x_n) dx_1 \ldots dx_n
\]

\[
= \frac{\text{var}(\bar{X}_n)}{\epsilon^2} = \frac{\sigma^2}{n \epsilon^2} \to 0 \quad \text{as} \quad n \to \infty.
\]

This completes the proof of the WLLN. The first inequality is justified because \(|\bar{x}_n - \mu| \geq \epsilon\), and therefore, \(1 \leq \frac{|\bar{x}_n - \mu|}{\epsilon}\), and consequently (squaring both sides), \(1 \leq \frac{|\bar{x}_n - \mu|^2}{\epsilon^2}\). The second equality is justified because the integral is everywhere non-negative. Therefore, extending the region of integration from the restricted set \( \{|x_1,\ldots,x_n|:|\bar{x}_n - \mu| \geq \epsilon\} \) to the whole space \( \mathbb{R}^n \) leads to an upper bound. This proof assumes the existence of the variance \( \sigma^2 \) of \( X_i \).

(b) There are many possible solutions here. Here is mine. I used this code in MATLAB to generate the random numbers and required plots:

```matlab
m=10000; r=10000;
X=10*randn([m,r])+100;
for j=1:r,
    Xn(j)=(1/j)*squeeze(sum(X(1:j,j),1));
end;
figure;plot(Xn);
figure;hist(X(100,1:100),20);
figure;hist(X(10000,1:10000),50);
```

Here we generated random variables \( \sim N(100, 100) \). The first plot (Fig. 2.15) illustrates the LLN because the arithmetic averages are shown to converge to the true mean (100) as the number of terms in the sum increases. The
2.30. Problems

histograms show that with only a few terms, we do not get a nice Gaussian, whereas using 10,000 terms, we get a nice bell curve. (If you chose a distribution other than normal, these histograms should reflect the chosen distribution.)

Problem 29. Derive the probability distribution of a biased random walk (i.e. let $p_\sigma = 1/2 + \delta$ and $p_{-\sigma} = 1/2 - \delta$ for some bias $0 \leq \delta \leq 1/2$).

Solution. By the CLT, the distribution will be Gaussian, of course. The mean step size is $\mu_i = \mathbb{E}X_i = \sigma(p_\sigma - p_{-\sigma}) = \sigma(1/2 + \delta - 1/2 + \delta) = 2\sigma\delta$. So the total displacement

$$X_{tot} = X_1 + X_2 + \cdots + X_n$$

has expectation value

$$\mathbb{E}X_{tot} = 2\sigma\delta n,$$

instead of 0. (i.e. it “drifts” linearly with time at constant speed $2\sigma\delta$.)

The variance is $\text{var}(X_i) = \mathbb{E}(X_i - \mu_i)^2 = p_\sigma(\sigma - 2\sigma\delta)^2 + p_{-\sigma}(-\sigma - 2\sigma\delta)^2 = (1/2 + \delta)\sigma^2(1 - 2\delta)^2 + (1/2 - \delta)\sigma^2(1 + 2\delta)^2 = \sigma^2(1/2 + \delta)(1 - 4\delta + 4\delta^2) + \sigma^2(1/2 - \delta)(1 + 4\delta + 4\delta^2) = \sigma^2[1 - 4\delta^2]$. The total variance is:

$$\text{var}(X_{tot}) = \text{var}(X_1) + \cdots + \text{var}(X_n) = \sigma^2[1 - 4\delta^2]n,$$

as opposed to $\sigma^2 n$. Thus, the variance is reduced. When $\delta = \pm 1/2$ (meaning steps are always to the left, or always to the right), then the variance is zero because the path is no longer random, but instead becomes deterministic.

Problem 30. Prove that in 3D the mean square displacement is $6Dt$, and in the general case of $d$ dimensions, it is equal to $2dDt$ (a direct calculation of the $d$-dimensional integral requires the spherical volume element in $d$-dim, which includes some Gamma functions).
Solution. In 3D the mean square displacement is
\[
E(r(t)^2) = \frac{1}{(4\pi Dt)^{3/2}} \int_{\mathbb{R}^3} r^2 \exp \left( -\frac{r^2}{4Dt} \right) d^3r
\]
\[
= \frac{1}{(4\pi Dt)^{3/2}} \int_{\mathbb{R}^3} r^2 \exp \left( -\frac{r^2}{4Dt} \right) r^2 dr \, d\cos(\theta) \, d\phi
\]
\[
= \frac{4\pi}{(4\pi Dt)^{3/2}} \int_0^\infty r^4 \exp \left( -\frac{r^2}{4Dt} \right) dr
\]
To integrate this we use the famous result \( \int_{-\infty}^{\infty} e^{-ax^2} \, dx = \sqrt{\pi/a} \), differentiate wrt \( a \) twice:
\[
\int_{-\infty}^{\infty} x^4 e^{-ax^2} \, dx = \frac{3}{4} \sqrt{\pi a} - \frac{5}{2}
\]
\[
\frac{4\pi}{(4\pi Dt)^{3/2}} \int_0^\infty r^4 \exp \left( -\frac{r^2}{4Dt} \right) dr = \frac{4\pi}{(4\pi Dt)^{3/2}} \frac{3}{8} \sqrt{\pi(4Dt)^{5/2}} = 6Dt.
\]
In the general case of \( d \) dimensions, the mean square displacement is:
\[
E(r(t)^2) = E(x_1^2 + \cdots + x_d^2) = 2dDt.
\]
There is no \( d \)-dimensional integral needed here, as each \( x_i^2 \) contributes \( 2Dt \), and there are \( d \) such terms, for a total of \( 2dDt \). ■

Problem 31. Prove that for the Poisson distribution the mean and variance are both equal to the parameter of the distribution.

Solution. Proofs can be found here:
http://filestore.aqa.org.uk/subjects/AQA-MS03-W-2-SM.PDF

Problem 32. Prove that Poisson distribution converges to a Gaussian in the limit of large \( n \). However, obtain the coefficient of the exponential as well (the prefactor), making use of the slightly more accurate Stirling’s formula.

Solution. The calculation we did previously was:
\[
\frac{e^{-\bar{n}n}/n!}{n^n} = \exp \left\{ -\bar{n} - \log n! + n \log \bar{n} \right\}
\]
\[
= \exp \left\{ -\bar{n} - n \log n + n + n \log \bar{n} \right\}
\]
\[
= \exp \left\{ (n - \bar{n}) + n \log(\bar{n}/n) \right\}
\]
\[
= \exp \left\{ (n - \bar{n}) + n \log \left[ 1 + \left( \frac{\bar{n} - n}{n} \right) \right] \right\}
\]
\[
\approx \exp \left\{ -\frac{(\bar{n} - n)^2}{2n} \right\} \approx \exp \left\{ -\frac{(\bar{n} - n)^2}{2\bar{n}} \right\}
\]
The prefactor \( \frac{1}{\sqrt{2\pi n}} \) is recovered by using
\[
n! \approx \sqrt{2\pi n} \left( \frac{n}{e} \right)^n.
\]
Now, replace \( \frac{1}{\sqrt{2\pi n}} e^{-\frac{n^2}{2n}} \) by \( \frac{1}{\sqrt{2\pi n}} e^{-\frac{n^2}{2n}} e^{-n \log n + n} \). This gives the desired result with the correct prefactor.

**Problem 33.** Let \( X \sim N(0, 1) \). Generate iid random numbers on a computer (say, 10,000 numbers). Those are the different realizations of \( X \), i.e. \( x_1, x_2, \ldots, x_{10,000} \). Next, consider another random variable, \( Y \).

(a) Let \( Y = X + 1 \), so that we now have 10,000 pairs of points: \( (x_1, y_1) \), \( (x_2, y_2) \), \ldots, \( (x_{10,000}, y_{10,000}) \). Plot these 10,000 pairs \( \{(x_i, y_i)\} \) as dots on scatter plot. Fit a straight line. What slope do you get? From the data, calculate the sample correlation coefficient. Is \( Y \) correlated to \( X \)? Why?

(b) Let \( X \) be as previously defined. Let \( Z \) be distributed identically to \( X \), but independent of \( X \). Generate random numbers on a computer to obtain pairs \( \{(x_i, z_i)\} \) of random numbers. Define a new random variable \( Y = X + Z \). Is \( Y \) correlated to \( X \)? Why? (Plot \( XY \) pairs on a scatter plot, fit a straight line, calculate \( r_{X,Y} \)).

(c) Let \( X \) be as defined previously. Let \( Y \sim N(0, 1) \). Generate random numbers for \( X \) and \( Y \), and plot the resulting pairs \( \{(x_i, y_i)\} \) on a scatter plot. Are \( X \) and \( Y \) correlated? Why?

**Solution.** (a) \( Y \) is correlated to \( X \) \((r = 1)\). On a scatter plot, we should see a perfect straight line (no deviation from it).

(b) \( \text{cov}(X, Y) = \text{cov}(X, X + Z) = \text{cov}(X, X) + \text{cov}(X, Z) = \text{var}(X) = 1 \), hence \( r = 1 \). Here on a scatter plot there will be random deviations from a straight line due to \( Z \). However, fitting a straight line will still give a slope of 1.

(c) Totally uncorrelated, since \( X \) and \( Y \) are independent. (Scatter plot looks random.)

**Problem 34.** Choose a distribution we have *not* used in class. Fix (choose) the parameters of the distribution. Let \( X \) be a random variable distributed accordingly. Calculate the mean and variance of \( X \) analytically (i.e. using the distribution function). Use a computer to generate random numbers according to the distribution of \( X \). (How do you generate such random numbers?) Plot the histogram of \( X \), compare to the PDF or PMF of \( X \) (plot both on the same graph). Calculate numerically the mean and variance of \( X \) (using the random numbers you generated). Compare to the true values of mean and variance obtained from the PDF or PMF.

**Problem 35.** Let \( X \) be the result of rolling a die. Generate \( n \) random numbers on a computer and obtain the random sample \( X_1, X_2, \ldots, X_n \). Take the arithmetic average: \( \overline{X}_n \equiv \frac{1}{n} \sum_{i=1}^{n} X_i \). Plot \( \overline{X}_n \) versus \( n \). What do you conclude? What theorem does this exercise illustrate?
Solution. This illustrates the law of large numbers. See the article for the plot:

Problem 36. Let \( Y \) be a Poisson rv with parameter \( \lambda \). Prove that \( Y \) can be written as the sum

\[ Y = X_1 + X_2 + \cdots + X_n, \]

where \( X_i \) are independent identically distributed rv’s, also with the Poisson distribution. What should be the lambda parameter of the \( X_i \)?

Solution. Let’s do the case of two rv’s. Let \( Z = X + Y \) where \( X \) and \( Y \) are Poisson, with parameters \( \lambda \) and \( \mu \), respectively. Then the PMF of \( Z \) is:

\[ P(Z = z) = \sum_{x=0}^{z} \frac{e^{-\lambda} \lambda^x}{x!} \frac{e^{-\mu} \mu^z-x}{(z-x)!} = \frac{e^{-(\lambda+\mu)}}{z!} \sum_{x=0}^{z} \binom{z}{x} \lambda^x \mu^{z-x} \]

Thus, \( Z \) is also Poisson, but with mean \( \lambda + \mu \). This can be extended to \( n \) Poisson variables. Their sum will also be Poisson. If \( Y = X_1 + \cdots + X_n \) has parameter \( \lambda \), then each \( X_i \) must have parameter \( \lambda/n \).

Problem 37. Let \( X_1, \ldots, X_n \) be a sequence of independent random variables with CDF’s \( F_n \) (\( X_i \) has CDF \( F \), \( i = 1, \ldots, n \)). Let \( X \) be a random variable with CDF \( F \). The sequence \( X_n \) is said to converge in distribution if the CDF’s converge pointwise, i.e.,

\[ \lim_{n \to \infty} F_n(x) = F(x), \]

at all points \( x \) for which \( F \) is continuous.

(a) Show that convergence of the CDF’s also implies the PDF’s. i.e. let \( f_i \) be the PDF of independent rv’s \( X_i \) (\( i = 1, \ldots, n \)) and \( f \) be the PDF of \( X \). Prove that convergence of the CDF’s implies:

\[ \lim_{n \to \infty} f_n(x) = f(x) \]

for all \( x \).

(b) Prove that the sequence of independent rv’s \( X_i \sim N(1/n, 1) \) converges in distribution to a standard normal random variable.

Solution. Since

\[ \frac{1}{\sqrt{2\pi}} \exp \left( -\frac{1}{2} \left[ x - \frac{1}{n} \right]^2 \right) \to \exp \left( -\frac{x^2}{2} \right), \]

it follows that \( X_n \) converges in distribution to \( X \sim N(0, 1) \).

Problem 38. Let \( X_n \sim N(0, 1/n) \) and let \( X = 0 \). Prove that for any \( \epsilon > 0 \),

\[ P(|X_n| > \epsilon) \to 0 \]
as \( n \to \infty \). This is an example of convergence in probability, i.e. \( P(|X_n| > \epsilon) \to 0 \) implies that \( X_n \) converges in probability to \( X (=0) \), since \( P(|X_n| > \epsilon) = P(|X_n - X| > \epsilon) \).

**Solution.** First we note that \( P(|X_n| > \epsilon) = P(|X_n|^2 > \epsilon^2) \). The latter is the integral:

\[
\int \{x_n^2 > \epsilon^2\} p(x_n)dx_n \leq \int \{x_n^2 > \epsilon^2\} \frac{x_n^2}{\epsilon^2} p(x_n)dx_n \leq \int_\mathbb{R} \frac{x_n^2}{\epsilon^2} p(x_n)dx_n
\]

\[
= \frac{\mathbb{E}X_n^2}{\epsilon^2} = \frac{\text{var}(X_n)}{\epsilon^2} = \frac{1}{n\epsilon^2} \to 0
\]

**Problem 39.** Let \( X_i \) be iidrv with uniform distribution over the interval [0,1]. Take the sum \( S_n = X_1 + X_2 + \cdots + X_n \). Find the distribution of \( X_n \) analytically (i.e. find its CDF and PDF). Show numerically (i.e. by generating random numbers on a computer) the histogram of \( S_n \) from \( n = 1, 2, \ldots, 10 \). What do you conclude?

**Solution.** This is straightforward and will be left as an exercise (simply generate random numbers in MATLAB to construct \( S_n \), and plot using the `hist` function). \( S_1 \) has the uniform distribution. \( S_2 \) has the “tent” distribution. etc. whereas \( S_n \) for large \( n \) looks more and more Gaussian as \( n \) increases, thanks to the CLT. Convergence to a Gaussian is very fast and does not require \( n \) to be very large.

**Problem 40.** Suppose that \( X \) has a PDF, \( p(x) = \frac{1}{2} \sin(x) \), where \( x \in [0, \pi] \), and equals zero elsewhere. Calculate its mean and variance. Calculate its skewness and kurtosis. Compare skewness and kurtosis to those of a normal distribution (with same mean and variance).

**Solution.** We will do the first two moments (others are obtained similarly):

\[
\mathbb{E}(X) = \int_0^\pi \frac{1}{2} \sin(x)dx = \frac{\pi}{2}.
\]

\[
\mathbb{E}(X - \frac{\pi}{2})^2 = \int_0^\pi \frac{1}{2} \sin(x)(x - \frac{\pi}{2})^2dx = \frac{1}{4}(\pi^2 - 8).
\]

Those results can be obtained from WolframAlpha by typing:

integrate x*(1/2)*sin(x) from 0 to Pi
integrate ((x-Pi/2)^2)*(1/2)*sin(x) from 0 to Pi

**Problem 41.** Let \( X \sim N(0,1) \). What is the distribution of \( Y = X^3 + 5 \)?
Solution.
\[ P(Y < y) = P(X^3 + 5 < y) = P(X \leq \sqrt[3]{y-5}) = \int_{-\infty}^{\sqrt[3]{y-5}} \frac{1}{\sqrt{2\pi}} e^{-x^2/2} \, dx. \]

It is also ok to calculate its PDF by differentiating the above CDF with respect to \( y \), making use of the Leibniz formula.

**Problem 42.** Calculate the mean and the variance of a random variable \( X \) distributed according to the PDF:
\[ p(x) = \frac{\gamma}{(x-\mu)^2 + \gamma^2}. \]

**Solution.** For the mean we have an integral of the type (set \( \gamma = 1, \mu = 0 \) without loss of generality, since I analyze the “tail” of the function here):
\[ \int_{-\infty}^{\infty} \frac{1}{x^2 + 1} \, dx \]

When \( x \) is large, this integral behaves like \( \int 1/x \sim \log(x) \), which diverges with \( x \). Thus, the mean does not exist. For the variance, we have an integral of the type
\[ \int_{-\infty}^{\infty} \frac{1}{x^2 + 1} x^2 \, dx \sim \int \, dx \sim x \to \infty \]
which also diverges. Thus, it has no variance.

**Problem 43.** The probability of \( k \) successes in \( n \) trials is \( (k = 0, 1, \ldots, n, \ 0 \leq p \leq 1) \):
\[ \mathbb{P}(k \text{ successes}) = \binom{n}{k} p^k (1-p)^{n-k}. \]

\( B \) is a random variable distributed as such. Prove that \( B \) has mean \( np \) and variance \( np(1-p) \).

**Solution.** See [https://en.wikipedia.org/wiki/Binomial_distribution](https://en.wikipedia.org/wiki/Binomial_distribution)

**Problem 44.** Suppose that you have a string instrument (e.g., electric guitar) whose strings, when plucked, behave like oscillators. The potential energy of the string is modeled by an anharmonic oscillator which consists of the sum of quadratic and quartic terms:
\[ V(x) = ax^2 + bx^4. \quad a, b \text{ non-negative constants} \]

The potential \( V \) is transferred to kinetic energy, which is then measured by the guitar’s pick-up coils and sent to the amplifier. The noise statistics of \( V \) are important for the design of the guitar amplifier circuits.

If the position \( x \) (\( x \): extension of the center of the string from its equilibrium position) is measured experimentally using an interferometer whose
2.30. Problems

Instrument noise is known to be normally distributed with mean $\mu$ and variance $\sigma^2$, what would you expect the noise statistics of $V$ to look like? (i.e. find the probability distribution of $V$) Note: you can assume there are no temporal correlations in the noise.

(a) When $b = 0$ and $a$ is nonzero (no anharmonicity).

(b) When $a = 0$ and $b$ is nonzero (anharmonic part only).

**Solution.** (a) When $V = ax^2$, the probability of $V < v$, $P(V < v)$, is:

$$P(ax^2 < v) = P(-\sqrt{\frac{v}{a}} < x < \sqrt{\frac{v}{a}}) = \frac{1}{\sqrt{2\pi\sigma^2}} \int_{-\sqrt{\frac{v}{a}}}^{\sqrt{\frac{v}{a}}} e^{-(x-\mu)^2/2\sigma^2} dx.$$  

Differentiating with respect to $v$ gives the PDF, $p_v(v) = \frac{dP(V < v)}{dv}$:

$$\frac{e^{-(\sqrt{\frac{v}{a}}-\mu)^2/2\sigma^2}}{\sqrt{2\pi\sigma^2}} \cdot \frac{d}{dv} \sqrt{\frac{v}{a}} - \text{(lower limit)} = \frac{e^{-(\sqrt{\frac{v}{a}}-\mu)^2/2\sigma^2} + e^{-(\sqrt{\frac{v}{a}}+\mu)^2/2\sigma^2}}{\sqrt{8av\pi\sigma^2}}.$$

(b) When $V = bx^4$ the probability of $V < v$, $P(V < v)$, is (imaginary roots are discarded, because probabilities are non-negative quantities):

$$P(bx^4 < v) = P(-\frac{1}{4}(v/b) < x < \frac{1}{4}(v/b)) = \frac{1}{\sqrt{2\pi\sigma^2}} \int_{-(v/b)^{1/4}}^{(v/b)^{1/4}} e^{-(x-\mu)^2/2\sigma^2} dx.$$  

The PDF is obtained by differentiation, $p_v(v) = \frac{dP(V < v)}{dv}$:

$$\frac{e^{-(\frac{1}{4}(v/b)^{1/4}-\mu)^2/2\sigma^2}}{\sqrt{2\pi\sigma^2}} \cdot \frac{d}{dv} \frac{1}{4}(v/b)^{1/4} - \text{(lower limit)} = \frac{1}{v^{3/4}} e^{-(\frac{1}{4}(v/b)^{1/4}-\mu)^2/2\sigma^2} + e^{-(\frac{1}{4}(v/b)^{1/4}+\mu)^2/2\sigma^2}}{\sqrt{32\pi\sigma^2 b^{1/4}}}.$$
Chapter 3

Propagation of Errors

In many experiments we are required to measure basic physical quantities and derive other quantities from these measurements. The derived quantities are often obtained from mathematical formulas. If we know the error in the basic measurements, what is the error in the derived quantities? This is the topic of error propagation. We begin with the single-variable case and extend to the multi-variable case.

3.1. Single Variable Case

Bragg’s law gives the angles for coherent and incoherent scattering from a crystal lattice. When X-rays are incident on an atom, they induce the radiation of electromagnetic waves at the same frequency, but the angular distribution depends on the lattice parameters. Crystalline solids produce specific patterns of reflected X-rays. At certain specific wavelengths and incident angles, we get intense peaks of reflected radiation (Bragg peaks). This is explained by modeling the crystal as a set of discrete parallel planes separated by a constant parameter $d$ (Fig. 3.1).

\begin{figure}[h]
\centering
\includegraphics[width=0.5\textwidth]{bragg_planes.png}
\caption{Bragg planes in a crystal.}
\end{figure}
The incident X-ray radiation produces a Bragg peak if the reflections off the various planes interfered constructively (Fig. 3.2). The interference is constructive when the phase shift is a multiple of $2\pi$ (Fig. 3.2).

This condition can be expressed by Bragg’s law,

$$n\lambda = 2d \sin \theta$$

The case $n = 1$ is the first order peak. For structure determination, the value of $\lambda$ (wavelength of incident radiation) would be known, and we would measure $\theta$ and solve for $d$ the lattice constant.

Another possible experiment we could do is determine $\lambda$ from a known value of $d$ and measured $\theta$. Suppose we measure $\theta$ and compute $\lambda$ from the Bragg formula. This is shown graphically in Fig. 3.3, where the error in $\theta$ is denoted $\alpha_\theta$. From this, we see that the error in $\lambda$ is:

$$\alpha_\lambda = |f(\overline{\theta} + \alpha_\theta) - f(\overline{\theta})| \approx \alpha_\theta \left| \frac{df}{d\theta}_{\theta=\overline{\theta}} \right|,$$

where in the last equality this error was obtained from calculus, by approximating the function $f$ by its first derivative at the point $\overline{\theta}$.

This linear approximation is illustrated graphically in Fig. 3.4. This is the most common method that people use for propagating errors. Another way to think about it is to view $\theta$ as a random variable that is the sum of a mean value $\overline{\theta}$ and a deviation $\delta\theta$:

$$\theta = \overline{\theta} + \delta\theta.$$

Figure 3.2. Bragg’s law. Two cases are shown: constructive vs destructive interference.
3.1. Single Variable Case

Figure 3.3. Error propagation for Bragg’s law.

Figure 3.4. Method for error propagation in a single variable based on linear approximation.

With this decomposition, \( \bar{\theta} \) is deterministic and \( \delta \theta \) is random. Taylor expansion of \( f \) to first order gives:

\[
\delta f = f(\bar{\theta} + \delta \theta) - f(\bar{\theta}) \approx \delta \theta \cdot \left( \frac{df}{d\theta} \bigg|_{\theta=\bar{\theta}} \right).
\]

Here \( \delta f \) is a random variable because it is equal to \( f(\bar{\theta} + \delta \theta) - f(\bar{\theta}) \), a function of a random variable \( \delta \theta \). Because \( \delta f \) is random, and its value changes every time an experiment is run, we can’t immediately use this expression and call \( \delta f \) the “propagated error”. We must take the additional step of computing its variance \( var(\delta f) \), which can then be used to obtain the error (as say, the square root of the variance).
The Taylor approximation term, \( \delta\theta \cdot \frac{df}{d\theta} \bigg|_{\theta=\bar{\theta}} \), is a random variable because \( \delta\theta \) is a random variable. \( \frac{df}{d\theta} \bigg|_{\theta=\bar{\theta}} \), on the other hand, is a deterministic quantity because it is the derivative of a deterministic function \( f \) evaluated at a deterministic argument \( \bar{\theta} \) is deterministic by definition, because the “mean” is just a number, hence deterministic).

Take the variance of \( \delta f \) and apply the property \( \text{var}(aX) = a^2 \text{var}(X) \):

\[
\text{var}(\delta f) = \left( \frac{df}{d\theta} \right)^2 \cdot \text{var}(\delta\theta).
\]

Because the square root of the variance is the standard deviation, let us take the error bar \( \alpha_f \) to be the standard deviation \( \sqrt{\text{var}(\delta f)} \) and similarly for \( \alpha_\theta = \sqrt{\text{var}(\delta\theta)} \), writing:

\[
\alpha_f = \left| \frac{df}{d\theta} \right| \alpha_\theta.
\]

This formula is identical to the one derived in the previous section. As an exercise, the reader should derive all the formulas in Table 3.1.

### 3.1.1. EXAMPLE: Error in Cosine (Single Variable).

Suppose we want to know the uncertainty in a cosine, i.e. \( f(\theta) = \cos \theta \), where \( \theta \) is a measured quantity. The measured angle is:

\[ \theta = 20 \pm 3^\circ \]

Then,

\[ (\cos \theta)_{\text{best}} = \cos 20^\circ = 0.94. \]

and

\[ \alpha_{\cos \theta} = \left| \frac{d \cos \theta}{d\theta} \right| \alpha_\theta = \left| \sin \theta \right| \alpha_\theta \quad \text{ (angles in radians)} \]
with $\alpha_\theta = 3^\circ = 0.05 \text{ rad}$. Note: the formula $d \cos \theta / d \theta = -\sin \theta$ only holds if $\theta$ is in radians. Then,

$$\alpha_{\cos \theta} = (\sin 20^\circ) \times 0.05 = 0.34 \times 0.05 = 0.02$$

so we report:

$$\cos \theta = 0.94 \pm 0.02$$

### 3.2. Multi-Variable Case

To avoid references to angles, let us switch notation from $\theta$ to $x$. Consider a function $f(x)$ which depends on several variables $x = (x_1, x_2, \ldots, x_n)$. The variables $x$ refer to different physical quantities measured in the laboratory. For example, suppose we measure the length ($L$), width ($W$) and depth ($D$) of a rectangular box, each with their own error bars, and we want to know the volume of the box and its uncertainty. Since $V = L \times W \times D$, we will learn to derive the following result:

$$\left( \frac{\alpha_V}{V} \right) = \sqrt{\left( \frac{\alpha_L}{L} \right)^2 + \left( \frac{\alpha_W}{W} \right)^2 + \left( \frac{\alpha_D}{D} \right)^2},$$

where $\alpha_V$ is the error in $V$, and similarly for $\alpha_L$, $\alpha_W$ and $\alpha_D$. Another experiment could involve Ohm’s law ($V = IR$; $R$, resistance; $I$, current; $V$, voltage). Suppose we measure voltage and current across a device and use this information to obtain the impedance ($R$) of the device. The error in $R$, $\alpha_R$ is related to the errors in $I$ and $V$, $\alpha_V$ and $\alpha_I$, as follows:

$$\left( \frac{\alpha_R}{R} \right) = \sqrt{\left( \frac{\alpha_I}{I} \right)^2 + \left( \frac{\alpha_V}{V} \right)^2}.$$ 

If the variables are statistically independent, the formula for error propagation takes a particularly simple form. The graphical method is easily generalized as follows:

$$|\delta f| = |f(x + \delta x) - f(x)|$$

and taking $\delta_{x_i}$ to be the error bars $\alpha_{x_i}$ in $x_i$ (where $\alpha_{x_i} > 0$), $\alpha_f$ to be the error bar in $f$:

$$|\alpha_f| = |f(x + \bar{\delta x}) - f(x)|.$$ 

Example in 2D:

$$|\alpha_f| = |f(A + \alpha_A, B + \alpha_B) - f(A, B)|.$$ 

Thus, if you know $f$, $A, B, \alpha_A, \alpha_B$ you can calculate the error bar in $f$. In general, this method should be sufficient.
In high dimensional spaces, it is impossible to visualize the function to be approximated. We can instead use derivatives:

\[ |\delta f| = |f(\mathbf{x} + \mathbf{\delta x}) - f(\mathbf{x})| \approx \left| \sum_{i=1}^{n} \frac{\partial f}{\partial x_i} \delta x_i \right|. \]

If you know the derivatives of \( f \), you can use this formula. Consider a function that adds several independent random variables:

\[ f(x_1, x_2, \ldots, x_n) = x_1 + x_2 + \cdots + x_n. \]

In that case, all derivatives are equal to 1: \( \partial f/\partial x_i = 1 \) for all \( i = 1, \ldots, n \). Denoting the errors in the \( x_i \) as \( \alpha_i \) and the error in \( \delta f \) as \( \alpha_f \), we find (since \( \alpha_i > 0 \)):

\[ |\alpha_f| = |\alpha_1 + \alpha_2 + \cdots + \alpha_n| = |\alpha_1| + |\alpha_2| + \cdots + |\alpha_n|. \]

We can also use the upper bound:

\[ \left| \sum_{i=1}^{n} \frac{\partial f}{\partial x_i} \delta x_i \right| \leq \sum_{i=1}^{n} \left| \frac{\partial f}{\partial x_i} \right| |\delta x_i|, \]

as our error bar in \( f \). Then, denoting the error bars in \( x_i \) as \( \alpha_{x_i} \) (instead of \( \delta x_i \)):

\[ |\alpha_f| = \sum_{i=1}^{n} \left| \frac{\partial f}{\partial x_i} \right| \cdot |\alpha_{x_i}|. \]

### 3.2.1. Remark: Adding Two Quantities

Consider the case \( f(x_1, x_2) = x_1 + x_2 \). One way to get the error in \( f \) in terms of the errors in \( x_1 \) and \( x_2 \) is to add the two error bars:

\[ |\alpha_f| = |\alpha_{x_1}| + |\alpha_{x_2}|. \]

This formula is easily derived using the graphical method:

\[ |\alpha_f| = |f(x_1 + \alpha_{x_1}, x_2 + \alpha_{x_2}) - f(x_1, x_2)| \approx \left| \frac{\partial f}{\partial x_1} \alpha_{x_1} + \frac{\partial f}{\partial x_2} \alpha_{x_2} \right| = |\alpha_{x_1}| + |\alpha_{x_2}|, \]

where \( f = x_1 + x_2 \), \( \alpha_{x_1} > 0 \) and \( \alpha_{x_2} > 0 \). (Note: we are treating all variables here as deterministic.)

Another method for obtaining the error bars in \( f \) when \( f(x_1, x_2) = x_1 + x_2 \) is to add error bars quadratically: \( \alpha_f = \sqrt{(\alpha_{x_1})^2 + (\alpha_{x_2})^2} \). Both methods are valid; they simply report different information. The method of adding errors in quadrature yields a tighter error bound than adding the errors linearly, i.e.

\[ |\alpha_{x_1}|^2 + |\alpha_{x_2}|^2 \leq |\alpha_{x_1}|^2 + |\alpha_{x_2}|^2 + 2|\alpha_{x_1}||\alpha_{x_2}| = (|\alpha_{x_1}| + |\alpha_{x_2}|)^2. \]
3.2. Multi-Variable Case

You can also see this graphically from the Pythagoras theorem: the hypotenuse \( \sqrt{(\alpha_{x_1})^2 + (\alpha_{x_2})^2} \) is always shorter than (or equal to) the sum of the two remaining sides \( |\alpha_{x_1}| + |\alpha_{x_2}| \).

3.2.2. Derivative Method: Case of Statistically Independent Variables. As in the 1D case, we view the experimental measurement as a random variable \( x \) which is the sum of a mean value \( \bar{x} \) (deterministic quantity) and a deviation \( \delta x \) (a random quantity):

\[
x = \bar{x} + \delta x.
\]

The Taylor expansion of \( f(x) \) about the point \( \bar{x} \) to first order is:

\[
\delta f = f(\bar{x} + \delta x) - f(\bar{x}) = \sum_{i=1}^{n} \frac{\partial f}{\partial x_i} \delta x_i + O(|\delta x|^2),
\]

where the derivatives are evaluated at the point \( \bar{x} \). Since \( \delta x \) are random variables, \( \delta f \) is also a random variable which depends on random variables \( \delta x = (\delta x_1, \ldots, \delta x_n) \). The \( \delta x_i \) are deviations from the point of expansion \( \bar{x} \). Again, we view the \( \partial f/\partial x_i \) as deterministic quantities whereas the \( \delta x_i \) are random. Note: \( \delta f \) itself is not the error bar. It’s a random variable. Its value changes every time a new experiment is done. However, an error can be obtained from \( \delta f \) by taking the square root of its variance.

If the random variables \( \{\delta x_i\} \) are mutually independent, taking the variance of Eq. (3.1) gives:

\[
\text{var}(\delta f) = \sum_{i=1}^{n} \text{var}(\partial_i f \cdot \delta x_i) = \sum_{i=1}^{n} (\partial_i f)^2 \text{var}(\delta x_i).
\]

Noting that \( \delta x = x - \bar{x} \), and that \( \text{var}(\delta x_i) = \text{var}(x_i) \) since constants such as \( \bar{x} \) don’t affect the variance, we have:

\[
\frac{\text{var}(\delta f)}{(\alpha_f)^2} = \sum_{i=1}^{n} \frac{(\partial_i f)^2 \text{var}(x_i)}{(\alpha_{x_i})^2},
\]

or

\[
(\alpha_f)^2 = \sum_{i=1}^{n} (\partial_i f)^2 (\alpha_{x_i})^2. \tag{3.2}
\]

How is this related to the result of the previous section?

\[
\sum_{i=1}^{n} |\partial_i f|^2 \cdot |\alpha_{x_i}|^2 \leq \left( \sum_{i=1}^{n} |\partial_i f| \cdot |\alpha_{x_i}| \right)^2.
\]

\( l_2 \) norm squared \( l_1 \) norm squared
### Table 3.2. Formulae for error propagation in several variables.

<table>
<thead>
<tr>
<th>( f(x_1, x_2, \ldots) )</th>
<th>Error Propagation</th>
</tr>
</thead>
<tbody>
<tr>
<td>( x_1 \pm x_2 )</td>
<td>( \alpha_f = \sqrt{(\alpha_{x_1})^2 + (\alpha_{x_2})^2} )</td>
</tr>
<tr>
<td>( x_1 \cdot x_2 ) or ( x_1/x_2 )</td>
<td>( \alpha_f = \sqrt{(\frac{\alpha_{x_1}}{x_1})^2 + \left(\frac{\alpha_{x_2}}{x_2}\right)^2} )</td>
</tr>
<tr>
<td>( x^n_1 )</td>
<td>( \alpha_f = \max\left(\frac{\alpha_{x_1}}{x_1}, \frac{\alpha_{x_2}}{x_2}\right) )</td>
</tr>
<tr>
<td>( k \frac{x_1}{x_2} )</td>
<td>( \alpha_f = \sqrt{\left(\frac{\alpha_{x_1}}{x_1}\right)^2 + \left(\frac{\alpha_{x_2}}{x_2}\right)^2} )</td>
</tr>
<tr>
<td>( x_1 + x_2 - x_3 + x_4 )</td>
<td>( \alpha_f = \sqrt{(\alpha_{x_1})^2 + (\alpha_{x_2})^2 + (\alpha_{x_3})^2 + (\alpha_{x_4})^2} )</td>
</tr>
</tbody>
</table>

The \( l_2 \)-norm gives tighter bounds than the \( l_1 \) norm. Recall: the \( l_p \)-norm of a vector \( \vec{x} \) is:

\[
||\vec{x}||_p = (|x_1|^p + \cdots + |x_n|^p)^{1/p}.
\]

The reader should derive all the formulas in Table 3.2.

Let us work out the first two lines explicitly. When \( x_1 \pm x_2 \), the partials \( \partial x_1 f = 1 \) and \( \partial x_2 f = \pm 1 \). Thus,

\[
(\alpha_f)^2 = \left(\frac{\partial f}{\partial x_1}\right)^2 (\alpha_{x_1})^2 + \left(\frac{\partial f}{\partial x_2}\right)^2 (\alpha_{x_2})^2
= 1 \cdot (\alpha_{x_1})^2 + (\pm 1)^2 \cdot (\alpha_{x_2})^2
= (\alpha_{x_1})^2 + (\alpha_{x_2})^2.
\]

Hence,

\[
\alpha_f = \sqrt{(\alpha_{x_1})^2 + (\alpha_{x_2})^2}.
\]

When \( f = x_1 \cdot x_2 \), we have \( \partial x_1 f = x_2 \) and \( \partial x_2 f = x_1 \). Then,

\[
(\alpha_f)^2 = \left(\frac{\partial f}{\partial x_1}\right)^2 (\alpha_{x_1})^2 + \left(\frac{\partial f}{\partial x_2}\right)^2 (\alpha_{x_2})^2
= x_2^2 \cdot (\alpha_{x_1})^2 + x_1^2 \cdot (\alpha_{x_2})^2
\]

Dividing both sides by \( f^2 = x_1^2 x_2^2 \) gives the result that fractional errors add up in quadrature:

\[
\left(\frac{\alpha_f}{f}\right)^2 = \left(\frac{\alpha_{x_1}}{x_1}\right)^2 + \left(\frac{\alpha_{x_2}}{x_2}\right)^2.
\]

We also get the same result for \( f = x_1/x_2 \).

#### 3.2.3. Example: Snell’s Law (Two Variables).

Consider light that enters a medium. Let \( (i) \) denote the angle of incidence and \( (r) \) the angle of
3.2. Multi-Variable Case

Figure 3.5. Snell’s law. The angle of incidence (in radians) is denoted \( i \); the angle of refraction is denoted \( r \).

refraction (Fig. 3.5). These two angles are related by Snell’s law, through the index of refraction of the medium \( n \) (in this case, glass),

\[
\sin i = n \sin r.
\]

We ask what is the error in \( n \) given experimentally measured errors in \( i \) and \( r \).

### 3.2.3.1. EXAMPLE: Method 1, Using Propagation of Errors Formula

From the previous lecture, \((\alpha_f)^2 = \sum_{i=1}^{n}(\partial_i f)^2(\alpha_x)^2\), and \( n = \frac{\sin i}{\sin r} \), we have

\[
(\alpha_n)^2 = \left(\frac{\cos i}{\sin r}\right)^2(\alpha_i)^2 + \left(\frac{\sin i \cdot \cos r}{(\sin r)^2}\right)^2(\alpha_r)^2.
\]

Dividing throughout by \( n^2 \), or equivalently, multiplying by \( 1/n^2 = (\sin r/\sin i)^2 \):

\[
\left(\frac{\alpha_n}{n}\right)^2 = \left(\frac{\cos i}{\sin i}\right)^2(\alpha_i)^2 + \left(\frac{\cos r}{\sin r}\right)^2(\alpha_r)^2.
\]

### 3.2.4. EXAMPLE: Method 2, Using Table of Formulas

Using the result for \( A/B \) in the table from the last page of the previous lecture, \( \alpha_z/z = \sqrt{\alpha_A/A}^2 + (\alpha_B/B)^2 \), and \( n = \sin i/\sin r \) with \( A = \sin i \) and \( B = \sin r \), we have:

\[
\frac{\alpha_n}{n} = \sqrt{\left(\frac{\alpha_{\sin i}}{\sin i}\right)^2 + \left(\frac{\alpha_{\sin r}}{\sin r}\right)^2},
\]

where \( \alpha_{\sin i} = |\cos i|\alpha_i \) and \( \alpha_{\sin r} = |\cos r|\alpha_r \). Some example numbers are provided in Table 3.3, for the case where errors in angles are \( \pm 1^\circ \), or 0.02 rad.

### 3.2.5. EXAMPLE: LC Resonant Circuit (Two Variables)

Suppose that we have an LC circuit (Fig. 3.6). Its resonance frequency can be shown
### Table 3.3. Error propagation through Snell’s law.

<table>
<thead>
<tr>
<th>(i) (deg)</th>
<th>(r) (deg)</th>
<th>(\sin i)</th>
<th>(\sin r)</th>
<th>(n)</th>
<th>(\frac{\alpha_{\sin i}}{\sin i})</th>
<th>(\frac{\alpha_{\sin r}}{\sin r})</th>
<th>(\frac{\alpha_n}{n})</th>
</tr>
</thead>
<tbody>
<tr>
<td>20</td>
<td>13</td>
<td>0.342</td>
<td>0.225</td>
<td>1.52</td>
<td>5%</td>
<td>8%</td>
<td>9%</td>
</tr>
<tr>
<td>40</td>
<td>23.5</td>
<td>0.643</td>
<td>0.399</td>
<td>1.61</td>
<td>2%</td>
<td>4%</td>
<td>5%</td>
</tr>
</tbody>
</table>

Figure 3.6. Resonant circuit.

It is generally difficult to measure \(L\) with good accuracy. A better method is to measure the resonant frequency \(f_0\) (which can be done using a network analyzer) and the capacitance \(C\). \(L\) is then given by

\[
L = \frac{1}{(2\pi f_0)^2 C}. 
\]

What about the error in \(L\)?

\[
\frac{(\alpha_L)^2}{L} = \left(\frac{\partial L}{\partial f_0}\right)^2 \left(\frac{\alpha_{f_0}}{f_0}\right)^2 + \left(\frac{\partial L}{\partial C}\right)^2 \left(\frac{\alpha_C}{C}\right)^2
\]

where

\[
\left|\frac{\partial L}{\partial f_0}\right| = \frac{1}{(2\pi f_0)^2} = \frac{2L}{f_0}
\]

and

\[
\left|\frac{\partial L}{\partial C}\right| = \frac{1}{(2\pi f_0)^2 C} = \frac{L}{C}.
\]

The result is:

\[
\left(\frac{\alpha_L}{L}\right)^2 = 4 \left(\frac{\alpha_{f_0}}{f_0}\right)^2 + \left(\frac{\alpha_C}{C}\right)^2,
\]

which could also have been derived from the above look-up table formula with \(z = kA^n/B^m\) for which

\[
\frac{\alpha_z}{z} = \sqrt{\left(\frac{m\alpha_A}{A}\right)^2 + \left(\frac{m\alpha_B}{B}\right)^2}.
\]

---

1This is seen by equating the voltages across each lumped element, i.e. \(j\omega L = \frac{1}{j\omega C}\), where \(j = \sqrt{-1}\) denotes an imaginary number and \(\omega = 2\pi f_0\).
3.2.6. EXAMPLE: Generic (Two Variables). Let’s determine the error in \( q = x^2y - xy^2 \), where we have measured experimentally the quantities:

\[
\begin{align*}
  x &= 3.0 \pm 0.1 \\
  y &= 2.0 \pm 0.1
\end{align*}
\]

The “\( x_{\text{best}} \)” value, as usual, is obtained from the reported values:

\[
q_{\text{best}} = 3^2 \cdot 2 - 3 \cdot 2^2 = 6.0.
\]

For the error, we need the following two terms:

\[
\begin{align*}
  \left| \frac{\partial q}{\partial x} \right| \alpha_x &= |2xy - y^2| \alpha_x = |12 - 4| \cdot 0.1 = 0.8, \\
  \left| \frac{\partial q}{\partial y} \right| \alpha_y &= |x^2 - 2xy| \alpha_y = |9 - 12| \cdot 0.1 = 0.3.
\end{align*}
\]

Adding the errors in quadrature gives:

\[
\alpha_q = \sqrt{(0.8)^2 + (0.3)^2} = 0.9.
\]

Therefore, we report:

\[
q = 6.0 \pm 0.9.
\]

3.3. When Variables are Not Statistically Independent

Previously, we found the following formula for propagation of errors

\[
(\alpha_f)^2 = \sum_{i=1}^{n} (\partial_i f)^2 (\alpha_{x_i})^2,
\]

where \( f \) is a function of \( n \) variables, i.e. \( f = f(x_1, x_2, \ldots, x_n) \). \( \partial_i f \) denotes the partial derivative of \( f \) with respect to \( x_i \). This result relies on the assumption that the different random variables are statistically independent. However, if the random variables are not statistically independent, covariance

\[
cov(X, Y) = \overline{XY} - \overline{X} \cdot \overline{Y}
\]

enters the picture and an additional term is required to describe the interdependence between the variables. Before we begin, let us remark the following properties of the covariance, which easily follow from the definition of covariance:

\[
\begin{align*}
  cov(X, X) &= \text{var}(X) \\
  cov(X, Y) &= cov(Y, X) \quad \text{(symmetry)} \\
  cov(X + Y, Z) &= cov(X, Z) + cov(Y, Z) \quad \text{(linearity in the first argument)} \\
  cov(aX, Y) &= a \cdot cov(X, Y) \quad \text{(linearity in the first argument)}
\end{align*}
\]

from which you can derive

\[
\begin{align*}
  cov(aX+bY, Z) &= a \cdot cov(X, Z) + b \cdot cov(Y, Z) \quad \text{(linearity in the first argument)} \\
  cov(Z, aX+bY) &= a \cdot cov(Z, X) + b \cdot cov(Z, Y) \quad \text{(linearity in the second argument)}
\end{align*}
\]
Also useful is the relationship:
\[
cov(a+X, Y) = \left(a + \bar{X}\right)\bar{Y} - (a + \bar{X}) \cdot \bar{Y} = a\bar{Y} + \bar{X}\bar{Y} - a\bar{Y} - \bar{X} \cdot \bar{Y} = cov(X, Y),
\]
since additive constants don’t change anything. Slightly more obvious is:
\[
cov(a, Y) = a\bar{Y} - a \cdot \bar{Y} = a\bar{Y} - a\bar{Y} = 0.
\]
Since \(cov\) is linear in both arguments, we say that it is \textit{bilinear}. Let \(\{X_1, \ldots, X_n\}\) be \(n\) random variables. Bilinearity gives:
\[
var(\sum_{i=1}^{n} a_i X_i) = \sum_{i,j=1}^{n} a_i a_j cov(X_i, X_j) = \sum_{i=1}^{n} a_i^2 var(X_i) + 2 \sum_{i<j} a_i a_j cov(X_i, X_j).
\]
So taking the covariance of
\[
(3.4) \quad \delta f = f(x + \delta x) - f(x) = \sum_{i=1}^{n} \frac{\partial f}{\partial x_i} \delta x_i + O(|\delta x|^2)
\]
with itself to get the variance (the square root of which can be taken as the error bar):
\[
var(\delta f) = cov(\delta f, \delta f) = \left(\sum_{i=1}^{n} \frac{\partial f}{\partial x_i} \delta x_i, \sum_{j=1}^{n} \frac{\partial f}{\partial x_j} \delta x_j\right)
\]
\[
= \sum_{i=1}^{n} (\partial_i f)^2 var(\delta x_i) + 2 \sum_{i<j} (\partial_i f)(\partial_j f) cov(\delta x_i, \delta x_j).
\]
Note: we neglected terms of order \(O(|\delta x|^2)\) and higher.
Recall that \(\delta x = x - \bar{x}\) and because \(\bar{x}\) is simply an additive constant, we have that:
\[
cov(\delta x_i, \delta x_j) = cov(x_i - \bar{x}_i, x_j - \bar{x}_j)
\]
\[
= cov(x_i, x_j) - cov(x_i, \bar{x}_j) - cov(\bar{x}_i, x_j) + cov(\bar{x}_i, \bar{x}_j).
\]
Observe that \(cov(\bar{x}_i, \bar{x}_j) = 0\) since both arguments are constant. Also, both \(cov(x_i, \bar{x}_j)\) and \(cov(\bar{x}_i, x_j)\) vanish since one of the arguments is a constant. This leaves:
\[
cov(\delta x_i, \delta x_j) = cov(x_i, x_j).
\]
Thus,
\[
var(f) = \sum_{i=1}^{n} (\partial_i f)^2 var(x_i) + 2 \sum_{i<j} (\partial_i f)(\partial_j f) cov(x_i, x_j).
\]
3.4. Several Functions of Several Variables

Suppose that \( f \) is a vector-valued function, i.e.
\[
\vec{f} = \begin{pmatrix} f_1 \\ f_2 \\ \vdots \\ f_m \end{pmatrix},
\]
then, one such relationship holds for all \( m \) components \( f_i \):
\[
(\alpha f_i)^2 = \sum_{j=1}^{n} (\partial_j f_i)^2 (\alpha x_j)^2 + 2 \sum_{h<j} (\partial_h f_i)(\partial_j f_i) \text{cov}(x_h, x_j),
\]
or
\[
\text{var}(f_i) = \sum_{j=1}^{n} (\partial_j f_i)^2 \text{var}(x_j) + 2 \sum_{h<j} (\partial_h f_i)(\partial_j f_i) \text{cov}(x_h, x_j).
\]

Of course, then \( m = 1 \), this expression reduces to Eq. (3.5).

Now suppose that we have \( m \) functions \( f_1, f_2, \ldots, f_m \), each a function of \( n \) different random variables \( x_1, x_2, \ldots, x_n \). In the general case, the different \( f_k \) will be correlated with one another, even if the \( x_1, x_2, \ldots, x_n \) are uncorrelated. The variances of the \( f_k \) are given by Eq. (3.6), whereas the covariances are:
\[
\text{cov}(f_k, f_l) = (f_k - \bar{f}_k)(f_l - \bar{f}_l) = \delta_{k,l} \frac{\partial f_k}{\partial x_i} \delta_{x_i} = \sum_{i=1}^{n} \frac{\partial f_k}{\partial x_i} \delta_{x_i} \delta_{x_i} = \sum_{i,j=1}^{n} \frac{\partial f_k}{\partial x_i} \frac{\partial f_l}{\partial x_j} \delta_{x_i} \delta_{x_j},
\]
which leads to:

\[
(3.7) \quad \text{cov}(f_k, f_l) = \sum_{i=1}^{n} \sum_{j=1}^{n} \left( \frac{\partial f_k}{\partial x_i} \right) \left( \frac{\partial f_l}{\partial x_j} \right) \text{cov}(x_i, x_j).
\]

We made use of

\[
\delta f_k = f_k(\mathbf{x} + \delta \mathbf{x}) - f_k(\mathbf{x}) = \sum_{i=1}^{n} \frac{\partial f_k}{\partial x_i} \delta x_i + O(|\delta \mathbf{x}|^2),
\]

and neglected terms of order \(O(|\delta \mathbf{x}|^2)\) and higher. If we denote \(g_{kl} = \frac{\partial f_l}{\partial x_k}\) the elements of a matrix \(G\) and \(F\) and \(X\) the covariance matrices of \(\mathbf{f}\) and \(\mathbf{x}\), respectively, then:

\[
F = GXG^T,
\]

which can also be written as:

\[
(3.8) \quad \text{cov}(\mathbf{f}, \mathbf{f}) = G \cdot \text{cov}(\mathbf{x}, \mathbf{x}) \cdot G^T,
\]

where \(\text{cov}(\mathbf{f}, \mathbf{f}) = F\) and \(\text{cov}(\mathbf{x}, \mathbf{x}) = X\). This very simple formula encodes all there is to know about error propagation.

### 3.5. Examples Using Matrix Method

We look at a few examples of how to apply Eq. (3.8). Suppose that \(m = 1\) and \(f : \mathbb{R}^3 \rightarrow \mathbb{R}\) depends on three variables, \(f = f(x, y, z)\). The errors on \(x, y, z\) are \(\sigma_x, \sigma_y, \sigma_z\), respectively. Then,

\[
G = \begin{pmatrix} \frac{\partial f}{\partial x} & \frac{\partial f}{\partial y} & \frac{\partial f}{\partial z} \end{pmatrix}
\]

\[
F = \begin{pmatrix} \frac{\partial f}{\partial x} & \frac{\partial f}{\partial y} & \frac{\partial f}{\partial z} \end{pmatrix} \begin{pmatrix} \sigma_x^2 & 0 & 0 \\ 0 & \sigma_y^2 & 0 \\ 0 & 0 & \sigma_z^2 \end{pmatrix} \begin{pmatrix} \frac{\partial f}{\partial x} & \frac{\partial f}{\partial y} & \frac{\partial f}{\partial z} \end{pmatrix}^T.
\]

This gives:

\[
F \equiv \text{var}(f) = \left( \frac{\partial f}{\partial x} \right)^2 \sigma_x^2 + \left( \frac{\partial f}{\partial y} \right)^2 \sigma_y^2 + \left( \frac{\partial f}{\partial z} \right)^2 \sigma_z^2.
\]

Suppose instead that \(f : \mathbb{R}^3 \rightarrow \mathbb{R}^3\) describes a change of coordinates from cylindrical to Cartesian, i.e. \(x = r \cos \phi, y = r \sin \phi\) and \(z = z\). Suppose that the random variables \(r, \phi, z\) are uncorrelated. Then,

\[
G = \begin{pmatrix} \cos \phi & -r \sin \phi & 0 \\ \sin \phi & r \cos \phi & 0 \\ 0 & 0 & 1 \end{pmatrix}
\]
### 3.7. Multiplicative Systematic Errors

Suppose instead that the systematic errors are multiplicative:

\[ x' = \delta \cdot x. \]

\[ F = \begin{pmatrix}
\cos \phi & -r \sin \phi & 0 \\
\sin \phi & r \cos \phi & 0 \\
0 & 0 & 1
\end{pmatrix}
\begin{pmatrix}
\sigma_r^2 & 0 & 0 \\
0 & \sigma_\phi^2 & 0 \\
0 & 0 & \sigma_z^2
\end{pmatrix}
\begin{pmatrix}
\cos \phi & \sin \phi & 0 \\
-r \sin \phi & r \cos \phi & 0 \\
0 & 0 & 1
\end{pmatrix}
\]

Then, we conclude that:

\[
\begin{pmatrix}
\sigma_r^2 \cos \phi + \sigma_\phi^2 r^2 \sin^2 \phi & (\sigma_r^2 - \sigma_\phi^2 r^2) \sin \phi \cos \phi & 0 \\
(\sigma_r^2 - \sigma_\phi^2 r^2) \sin \phi \cos \phi & \sigma_r^2 \sin^2 \phi + \sigma_\phi^2 r^2 \cos^2 \phi & 0 \\
0 & 0 & \sigma_z^2
\end{pmatrix}
\]

We see that the errors and correlations involving \( z \) are not affected by the transformation (since \( z \) is never transformed). The errors in \( x \) and \( y \) depend on where they are \((r, \phi)\) and on the variances of \( r, \phi \).

### 3.6. Additive Systematic Errors

Systematic errors often appear as an overall shift of the value we measure, i.e. instead of measuring \( x \) we measure \( x' \) which is the sum of \( x \) and an offset \( \delta \):

\[ x' = x + \delta. \]

Another random variable \( y \) is subject to the same systematic drift:

\[ y' = y + \delta. \]

Thus, \( x \) and \( y \) have a common systematic error. Assume that \( x \) and \( y \) are independent of each other and independent of \( \delta \). The error in \( x' \) is:

\[ \text{var}(x') = \text{var}(x) + \text{var}(\delta) \]

and similarly for \( y \). The covariance is:

\[ \text{cov}(x', y') = \text{cov}(x + \delta, y + \delta) = \text{var}(\delta). \]

Thus, the covariance matrix for \( x', y' \) has the random and systematic error added in quadrature along the diagonal whereas off-diagonal elements are the square systematic errors:

\[
\begin{pmatrix}
\text{var}(x) & \text{cov}(x, y) \\
\text{cov}(y, x) & \text{var}(y)
\end{pmatrix}
= \begin{pmatrix}
\text{var}(x) + \text{var}(\delta) & \text{var}(\delta) \\
\text{var}(\delta) & \text{var}(y) + \text{var}(\delta)
\end{pmatrix}.
\]

If the systematic error is constant (i.e. does not fluctuate), then \( \text{var}(\delta) = 0 \) and \( x', y' \) are uncorrelated.

### 3.7. Multiplicative Systematic Errors

Suppose instead that the systematic errors are multiplicative:

\[ x' = \delta \cdot x. \]
Similarly for $y$:

$$y' = \delta \cdot y.$$ 

We assume that $x$ and $y$ are independent of each other. We also assume that $\delta$ is independent of $x$ and $y$. Then,

$$\text{var}(x') = \delta^2 \text{var}(x) + \overline{x}^2 \text{var}(\delta) + \text{var}(\delta) \text{var}(x)$$

and

$$\text{var}(y') = \delta^2 \text{var}(y) + \overline{y}^2 \text{var}(\delta) + \text{var}(\delta) \text{var}(y).$$

The covariances are:

$$\text{cov}(x', y') = (\delta \cdot x - \overline{x} \cdot \delta \cdot y - \overline{y} \cdot \delta \cdot x) = \delta \cdot x \delta \cdot y - \overline{x} \cdot \delta \cdot y = \text{var}(\delta) \overline{x} \cdot \overline{y}.$$

The covariance matrix of $x', y'$ is therefore:

$$\begin{pmatrix}
\delta^2 \text{var}(x) + \overline{x}^2 \text{var}(\delta) + \text{var}(\delta) \text{var}(x) & \text{var}(\delta) \overline{x} \cdot \overline{y} \\
\text{var}(\delta) \overline{x} \cdot \overline{y} & \delta^2 \text{var}(y) + \overline{y}^2 \text{var}(\delta) + \text{var}(\delta) \text{var}(y)
\end{pmatrix}.$$ 

The amount of fluctuations in $\delta$ determine the magnitude of covariance between $x'$ and $y'$. If $\delta$ does not fluctuate, $\text{var}(\delta) = 0$, $\delta$ is just a number and this reduces to:

$$\begin{pmatrix}
\delta^2 \text{var}(x) & 0 \\
0 & \delta^2 \text{var}(y)
\end{pmatrix}.$$ 

There is no longer any covariance between $x'$ and $y'$. As expected, the errors in $x', y'$ depend on the mean value of the scaling factor $\delta$. Since the multiplicative factor is just a constant, this result could also have been deduced from the theorem $\text{var}(aX) = a^2 \text{var}(X)$.

### 3.8. Problems

**Problem 45.** The Richter magnitude of an earthquake is determined from the logarithm of the amplitude of waves recorded by seismographs (adjustments are included to compensate for the variation in the distance between the various seismographs and the epicenter of the earthquake). The formula is:

$$M_L = \log_{10}(A/A_0(\delta))$$

where $A$ is the maximum excursion of the Wood-Anderson seismograph, the empirical function $A_0$ depends only on the epicentral distance of the station, $\delta$. Both $A$ and $\delta$ are prone to measurement error. Find the uncertainty in $M_L$ due to errors in $A$ and $\delta$.

**Solution.** Propagation of error formula is:

$$\sigma_{M_L}^2 = \left| \frac{\partial M_L}{\partial A} \right|^2 \sigma_A^2 + \left| \frac{\partial M_L}{\partial \delta} \right|^2 \sigma_\delta^2.$$
where the derivatives are:

\[
\begin{align*}
\frac{\partial M_L}{\partial A} &= \frac{1}{\log(10)} \frac{A_0(\delta)}{A} \cdot \frac{1}{A_0(\delta)} = \frac{1}{A \cdot \log(10)}, \\
\frac{\partial M_L}{\partial \delta} &= -\frac{\partial A_0(\delta) / \partial \delta}{A_0(\delta) \cdot \log(10)}. 
\end{align*}
\]

**Problem 46.** You built a setup to detect light with an avalanche photodiode. In an avalanche photodiode the carrier production rate \( \eta(P/h\nu) \) is increased by a factor of \( M \) because of the ionization by the drifting electrons and holes. Here, \( \eta \) is the detector quantum efficiency (number of carriers generated divided by the number of photons absorbed), \( P \) is the power absorbed in the detector and \( h\nu \) is the incident photon energy. The photocurrent is enhanced by the same value:

\[
i = Me \left[ \eta \left( \frac{P}{h\nu} \right) + g \right]
\]

where \( g \) is a constant and \( e \) is the charge of the electron. Suppose that light incident on the photodetector is in a coherent state. A coherent state is a photon field that is a linear superposition of single-mode states. It is the closest approximation to a classical electromagnetic field, as far as quantum mechanics’ Heisenberg uncertainty principle allows.

In a coherent state the fluctuations of the photon number follow Poisson statistics, i.e. for a measurement of the number of photons in the light field, the probability of the field containing \( n \) photons is

\[
P(n; \bar{n}) = \frac{e^{-\bar{n}} \bar{n}^n}{n!}
\]

which is a Poisson distribution with mean \( \bar{n} \). The light power absorbed in the detector \( (P) \) is proportional to \( n \), i.e. \( P = Cn \) where \( C \) is a constant.

(a) Compute the fluctuations in the photon count \( (n) \) of the coherent state in terms of the measured photocurrent \( (i) \). From the noise, estimate the actual photon count \( n \) in the electromagnetic field.

(b) Same as (a) but use the fluctuations in the measured electrical power instead of current \( (i) \). Electrical power is \( i^2R \), where \( R \) is a constant (the load resistance).

**Solution.** (a)

\[
i = Me \left[ \eta \frac{Cn}{h\nu} + g \right] \rightarrow \delta i = \left| M \eta \frac{C}{h\nu} \right| \delta n \rightarrow \delta n = \frac{\delta i}{\left| M \eta \frac{C}{h\nu} \right|}
\]
square $\delta n$ to get $n$.

(b) 

\[ P = i^2 R = R(Me)^2 \left( \frac{Cn}{h\nu} + g \right)^2 \]

expanding the square bracket:

\[ \eta^2 \frac{c^2 n^2}{(h\nu)^2} + 2\eta \frac{Cng}{h\nu} + g^2 \]

solving for $\delta P$

\[ \delta P = R(Me)^2 \left[ 2\eta^2 \frac{c^2 n}{(h\nu)^2} + 2\eta Cg \right] \delta n \]

replace $n$ by $(\delta n)^2$ and solve to get $\delta n$. Square to get $n$. ■

**Problem 47.** Use the method of error propagation to find the error in the work function $W$ (due to a force $F(x)$ applied over a distance $x$, i.e. $W = \int F(x)dx$ and you can assume Hooke’s law for $F$) performed by dragging a light object of mass $m$ across a distance $x$. The object is mounted at the end of a spring whose constant is $k$. The error in $x$ is denoted $\delta x$.

(a) Find the error $\delta W$ in terms of $\delta x$.

(b) What if the displacement $x$ is measured using an interferometer whose output is designed to produce a voltage $V$ (you may assume that $V$ is linear in $x$).

**Solution.** (a) Substitute $F = kx$ and integrate to get $W = \frac{1}{2}kx^2$. Then $\delta W = |\partial W/\partial x| \delta x = |kx| \delta x$.

(b) Now we have $x = AV$, where $A$ is some constant. Then, $W = \frac{1}{2}kA^2 V^2$ and $\delta W = |kA^2 V| \delta V$. ■

**Problem 48.** We have seen in class that when two random variables $X$ and $Y$ are added to form a new random variable, $Z = X + Y$, the errors add in quadrature:

\[ \sigma_Z^2 = \sigma_X^2 + \sigma_Y^2. \]

However, in your chemistry labs, you may have been taught that the errors add in first power:

\[ \sigma_Z = \sigma_X + \sigma_Y. \]

There are no restrictions on the distributions of $X$ and $Y$ apart from the obvious requirement of finite variances.

(a) Derive the second formula ($\sigma_Z = \sigma_X + \sigma_Y$, where $\sigma_X, \sigma_Y \geq 0$).
(b) Show the relationship between these two different methods for computing the error in \( Z \), i.e., derive an inequality that relates these two different forms for \( \sigma_Z \).

**Solution.** (a) The second formula follows from mapping the range of \( X \) and \( Y \) values onto the \( Z \) axis. \( Z(X,Y) = X + Y \) is a function of two variables. As \( X \) and \( Y \) range over their allowed values (thanks to \( \sigma_X \) and \( \sigma_Y \)), \( Z \) ranges from

\[
Z_{\text{min}} = (X - \sigma_X) + (Y - \sigma_Y)
\]

up to

\[
Z_{\text{max}} = (X + \sigma_X) + (Y + \sigma_Y)
\]

The difference between \( Z_{\text{min}} \) and \( Z_{\text{max}} \) is:

\[
Z_{\text{max}} - Z_{\text{min}} = 2\sigma_X + 2\sigma_Y.
\]

This is the total range. Error bar is half of this, since we write it symmetrically as \( Z \pm \delta Z \):

\[
\sigma_Z = \sigma_X + \sigma_Y.
\]

(b) Observe that \( (\sigma_X, \sigma_Y \geq 0) \):

\[
|\sigma_X + \sigma_Y|^2 = \sigma_X^2 + \sigma_Y^2 + 2\sigma_X\sigma_Y.
\]

Hence,

\[
|\sigma_X + \sigma_Y|^2 = \sigma_X^2 + \sigma_Y^2 + 2\sigma_X\sigma_Y \geq \sigma_Z^2 + \sigma_Y^2.
\]

The left hand side is \( \sigma_Z^2 \), where \( \sigma_Z = \sigma_X + \sigma_Y \). The right hand side is \( \sigma_Y^2 \), where \( \sigma_Z^2 = \sigma_X^2 + \sigma_Y^2 \). Thus, the error bar obtained from \( \sigma_Z = \sigma_X + \sigma_Y \) is larger than the error bars obtained by adding the errors in quadrature. ■

**Problem 49.** Use error propagation to show how an error in the measured radius leads to errors in the calculation of sphere volume (as calculated from the radius).

**Solution.** From \( V = \frac{4}{3}\pi r^3 \), \( \delta V = 4\pi r^2 \delta r \). ■

**Problem 50.** Same question as the previous one, except that now consider the case where you only remember the value of \( \pi \) to a few digits (e.g. \( \pi = 3.14 \)). How does this error in \( \pi \) (in addition to the error in \( r \)) contribute to the uncertainty in volume?

**Solution.** \( (\delta V)^2 = |4\pi r^2|^2 |\delta r|^2 + \left(\frac{4}{3}r^3\right)^2 (|\delta \pi|^2) \). ■

**Problem 51.** Suppose you work at a particle accelerator and your job is to spend your life measuring two quantities in the laboratory, \( X \) and \( Y \). \( X \) could be for example, a voltage, whereas \( Y \) could be a position. Both \( X \) and \( Y \) relate to some important physical quantity being measured. The two measurements are uncorrelated and statistically independent. It
is known that the values of $X$ range continuously from 0 to 1 and appear to be uniformly distributed whereas $Y$ appears to be Gaussian distributed with mean 0 and variance 1.

(a) Calculate the value of the probability $P(0.2 < X < 0.8, Y < 0)$.
(b) Calculate the probability $P(0.2 < X < 0.8, Y \neq 0)$ (i.e. the joint probability that $X$ lies in the range $[0.2, 0.8]$ and $Y$ does not equal zero).
(c) A physical theory for the new particle predicts that its spin can be obtained from $Z = Xe^Y$, where $X$ is related to the total orbital moment of the subatomic particles and $Y$ is related to the total spin of these particles. Definite values of $Z$ are hard to obtain due to experimental error, so the best you can do is set an upper bound on its value. Explain how you would set such an upper bound. Or equivalently, calculate the chance $P(Z < z)$ that $Z$ will take a value less than $z$.

Solution. (a) $P = 0.6 \times 0.5 = 0.30$
(b) $P = 0.6 \times 1 = 0.6$
(c) Method 1: since the distributions of $X$ and $Y$ are known:

$$P(Z < z) = \mathbb{P}(Xe^Y < z) = \int_{\{(x,y)|xe^y<z,0\leq x\leq 1\}} p_{XY}(x,y)dx\,dy$$

$$= \int_{\{(x,y)|xe^y<z,0\leq x\leq 1\}} \frac{1}{\sqrt{2\pi}} e^{-y^2/2} \, dx\,dy = \int_0^1 dx \int_{-\infty}^{\log(z/x)} \frac{e^{-y^2/2}}{\sqrt{2\pi}} \, dy$$

$$= \int_0^1 \Phi(\log(z/x)) \, dx.$$

Method 2: using error propagation, which can be done since $z(x,y)$ is known and the errors in $X$ and $Y$ are also known.

Problem 52. From the coast of Normandy, you measure the position of an English flag along the coast of Great Britain, across the English channel, using a telescope. You use the angle the telescope makes relative to its base, and the known distance across the English channel, to convert the angle to lateral distance along the coast (using simple trigonometry). However, the intense winds make it difficult if not nearly impossible for you to readout the angle (and hence the distance) properly. The intense vibrations result in random fluctuations of this distance.

Fortunately, you know that averaging several measurements together can reduce the noise. Let $X_1, \ldots, X_n$ be $n$ measurements of the flag’s true position $X$ (r.v. mean $\mu$ and variance $\sigma^2$; the variance is a measure of the area of the flag). What should the value of $n$ be ($n$: no. of samples acquired in an experiment) needed to ensure that the probability that the position
(calculated form the sample mean) does not deviate from the true position of the flag by more than $\sigma/10$ is at least 0.95.

**Solution.** Using the inequality on p.1 with $\epsilon = \sigma/10$

$$P\left(\left|\bar{X}_n - \mu\right| \leq \frac{\sigma}{10}\right) = 1 - P\left(\left|\bar{X}_n - \mu\right| > \frac{\sigma}{10}\right) \geq 1 - \frac{\text{var}(\bar{X}_n)}{(\sigma/10)^2}$$

Where $\bar{X}_n = \frac{X_1 + \cdots + X_n}{n}$, $\text{var}(\bar{X}_n) = \frac{\text{var}(X)}{n} = \frac{\sigma^2}{n}$.

$$P\left(\left|\bar{X}_n - \mu\right| \leq \frac{\sigma}{10}\right) \geq 1 - \frac{\text{var}(X)}{n(\sigma/10)^2} = 1 - \frac{100}{n}.$$

If we want this probability to be at least 0.95, we must have $100/n \leq 0.05$ or $n \geq 100/0.05 = 2000.$ □

**Problem 53.** Suppose we have Hooke’s law $F = kx$, where $k$ and $x$ have uncertainties. Use error propagation to determine the error in $F$ from the errors in $k$ and $x$.

**Solution.**

$$(\delta F)^2 = \left(\frac{\partial F}{\partial k}\right)^2 (\delta k)^2 + \left(\frac{\partial F}{\partial x}\right)^2 (\delta x)^2 = x^2(\delta k)^2 + k^2(\delta x)^2$$

$$(\delta F) = \sqrt{x^2(\delta k)^2 + k^2(\delta x)^2}.$$ □
Statistical Parameter Estimation

Statistical distributions, such as the Gaussian distribution, contain parameters such as the mean and variance whose values may be unknown. For example, suppose that we want to estimate the mean of a distribution. We may have at our disposal a series of measurements of a random variable \( X \) (random sample) \( \{X_1, \ldots, X_n\} \). How should we calculate the average?

For example, should we use the arithmetic mean (sample mean),

\[
\langle X \rangle_{\text{arith.}} \equiv \frac{1}{n} \sum_{i=1}^{n} X_i,
\]

the harmonic mean

\[
\langle X \rangle_{\text{harm.}} \equiv \left( \frac{1}{n} \sum_{i=1}^{n} X_i^{-1} \right)^{-1},
\]

or the geometric mean

\[
\langle X \rangle_{\text{geom.}} \equiv \left( \prod_{i=1}^{n} X_i \right)^{1/n}.
\]

Methods have been developed to estimate the parameters of statistical distributions. Estimation theory is a branch of statistics that deals with estimating the values of parameters based on measured empirical data. An estimator is not an exact representation of the parameters, but instead attempts to approximate the parameters from the measurements using some
underlying principle. There are several estimation methods that are commonly used, such as maximum likelihood estimation (MLE), Bayes estimation, the method of moments, the maximum a posteriori (MAP) method, etc. The simplest one is MLE; other methods require additional theory. We shall cover the MLE method.

4.1. Maximum Likelihood Estimation (MLE)

MLE estimates the parameters of a statistical model by finding the parameter values that maximize the likelihood of making the observations given the parameters.

4.1.1. Likelihood Function. Let us construct a so-called “likelihood function” \( L(\theta) \), as a function of the unknown parameters \( \theta = (\theta_1, \ldots, \theta_p) \), and conditioned by the observation of the random samples. We will set this likelihood function to be equal to the joint probability density of observing the values \( X_1, X_2, \ldots, X_n \). Since the random samples are independent random variables, the joint density factorizes into a product of densities for each:

\[
L(\theta|X_1, X_2, \ldots, X_n) \equiv p_\theta(X_1, X_2, \ldots, X_n) = \prod_{i=1}^{n} p_{X_i}(X_i|\theta).
\]

A simple interpretation of the likelihood function is as follows: if \( x_1, x_2, \ldots, x_n \) denote the values taken by the random variables \( X_1, X_2, \ldots, X_n \), and supposed that these rv’s are iidrv, then:

\[
\text{probability that } x_i \text{ is in } [x_i, x_i + dx_i] \text{ for all } i = \prod_{i=1}^{n} p(x_i|\theta)dx_i,
\]

where \( \theta \) represents one or more parameters of the distribution \( p(x|\theta) \). The likelihood function, \( L(\theta) = \prod_{i=1}^{n} p(x_i|\theta) \), which is just the joint pdf of the \( x_i \), is treated here as a function of the parameter(s), \( \theta \). The \( x_i \), on the other hand, are treated as fixed (the experiment is over). The MLE method consists of solving the system of equations, \( \nabla_\theta L = 0 \), or \( \partial L/\partial \theta_i = 0, i = 1, \ldots, p \) for the unknown parameters \( \theta_i \), as function of the data, \( x_1, \ldots, x_n \).

Next, we need to know or assume a distribution for the random variables \( X_1, X_2, \ldots, X_n \). For example, if they are normally distributed, the parameters are \( \theta = (\sigma, \mu) \). Due to statistical independence, the joint probability distribution \( p_{\mu,\sigma}(X_1, X_2, \ldots, X_n) \) for observing the whole set of \( n \) readouts \( \{X_i\} \) factorizes as a product:

\[
p_{X_1}(X_1) \cdot p_{X_2}(X_2) \cdots p_{X_n}(X_n) = \frac{1}{(2\pi)^{n/2}\sigma^n} e^{-\sum_{i=1}^{n}(X_i-\mu)^2/2\sigma^2}.
\]
The problem we want to solve is this: given the measurements \( \{X_1, \ldots, X_n\} \) we want to estimate the unknown quantities \( \mu \) and \( \sigma \). In other words, we would like to find “best estimates” for these parameters in terms of the observations \( X_1, \ldots, X_n \).

### 4.1.2. Principle of Maximum Likelihood

The principle of maximum likelihood states that the best estimates for \( \theta \) are those for which the observed data \( X_1, \ldots, X_n \) are most likely to occur, i.e., for which the likelihood function \( L(\theta | X_1, X_2, \ldots, X_n) \) is a maximum with respect to \( \theta \). Suppose that \( \theta = (\mu, \sigma) \), maximization with respect to \( \theta \) means that we should enforce the following two conditions:\(^1\)

\[
\frac{\partial L}{\partial \mu} = 0 \quad \text{and} \quad \frac{\partial L}{\partial \sigma} = 0,
\]

where the derivatives are evaluated at the point \( (\mu, \sigma) = (\hat{\mu}, \hat{\sigma}) \). The hat notation is used to denote the particular choice of \( \theta \) values obtained by MLE.

Let us start with the first condition \( \left( \frac{\partial L}{\partial \mu} = 0 \right) \). Inspection of the above “Gaussian” expression for \( p_{\mu, \sigma} \) shows that \( \mu \) only appears in the argument of exp. Finding the extremum of \( \exp(f(\mu)) \) with respect to \( \mu \) is equivalent to finding the extremum of \( f(\mu) \) with respect to \( \mu \), since:

\[
\frac{\partial}{\partial \mu} e^{f(\mu)} = e^{f(\mu)} \frac{\partial f}{\partial \mu} = 0
\]

and dividing both sides by \( e^{f(\mu)} \) implies that \( \frac{\partial f}{\partial \mu} = 0 \). In general, you can find the extremum of \( \log(L) \) rather than \( L \); this is sometimes easier.\(^2\) Thus,

\[
\frac{\partial}{\partial \mu} \sum_{i=1}^{n} \frac{(X_i - \mu)^2}{2\sigma^2} = 0 \quad \text{or} \quad \sum_{i=1}^{n} (X_i - \mu) = 0,
\]

---

1. An extremum of \( L \) is found by setting \( dL = 0 \). Take \( \theta = (\mu, \sigma) \), for example: since \( dL(\mu, \sigma) = \partial_\mu dL + \partial_\sigma dL \) and \( d\mu, d\sigma \) are arbitrary displacements, \( dL = 0 \) implies that the gradient of \( L \) vanishes: \( \nabla L(\mu, \sigma) = 0 \). The vanishing gradient means that all partial derivatives vanish: \( \partial_\mu = 0 \) and \( \partial_\sigma = 0 \).

2. We note that finding the extremum of \( L \) is equivalent to finding the extremum of \( \log L \). This is because \( \log(x) \) function is a monotonic function. Indeed, a necessary condition for the maximization of \( \log L \) is:

\[
\frac{\partial}{\partial x} \log L(x) = \frac{1}{L(x)} \frac{\partial}{\partial x} L(x) = 0.
\]

Multiplying both sides of the equation by \( L(x) \) leads to \( \frac{\partial}{\partial x} L(x) = 0 \), which is a necessary condition for a maximum in \( L(x) \). Thus, maximizing \( L \) is the same as maximizing \( \log L \). (Exercise: our proof is not complete since we have only considered the necessary condition for a maximum. Can you complete the argument by analyzing the sufficient condition?)
which leads to the formula for sample mean we have been using all along:

$$\hat{\mu} = \frac{1}{n} \sum_{i=1}^{n} X_i.$$ 

We shall write $\hat{\mu}$ with a hat to denote that this particular value of $\mu$ corresponds to the MLE for the mean. This provides a justification for the use of $\hat{\mu}$ as the value of $x_{best}$.

For $\partial L / \partial \sigma$ the situation is different because $\sigma$ also appears in the coefficient of $\exp$. In this case, we look for the extremum of $L$ instead of its log. Differentiating with respect to $\sigma$ we get

$$(2\pi)^{n/2} \frac{\partial L}{\partial \sigma} = \left( -n \right) e^{-\Sigma_i} + \frac{1}{\sigma^n} e^{-\Sigma_i} \left( - \sum_{i=1}^{n} \frac{(X_i - \mu)^2}{2} \right) \left( - \frac{2}{\sigma^3} \right) = 0.$$

Divide by $e^{-\Sigma_i}$, multiply by $\sigma^{n+1}$:

$$-n + \frac{1}{\sigma^2} \left( \sum_{i=1}^{n} (X_i - \mu)^2 \right) = 0$$

which gives

$$\hat{\sigma}^2_n = \frac{1}{n} \sum_{i=1}^{n} (X_i - \hat{\mu})^2$$

where $\hat{\mu}$ is the sample mean and $\Sigma_i$ is shorthand for $\sum_{i=1}^{n} (X_i - \mu)^2 / 2\sigma^2$. In the last step we wrote $(\hat{\mu}, \hat{\sigma})$ to denote the fact that this particular choice of values for $(\mu, \sigma)$ corresponds to the MLE.

In practice, we instead use the quantity with the coefficient $1/(n-1)$ instead of $1/n$:

$$\hat{\sigma}^2_{n-1} = \frac{1}{n-1} \sum_{i=1}^{n} (X_i - \hat{\mu})^2$$

sample variance

because the $n-1$ corrects for the underestimation of $\sigma^2$ when we use $\hat{\mu}$ (sample mean) as our estimator for $\mu$. Recall that the choice of $\mu = \hat{\mu}$ (sample mean), by definition, minimizes the quantity $\sum_i (X_i - \mu)^2$ for $\mu = \hat{\mu}$. We say that $\sigma^2$ is a biased estimator of the variance whereas $\hat{\sigma}^2_{n-1}$ is an unbiased estimator. The proof for the $n/(n-1)$ correction factor is found in Section 4.4.

4.2. Biased vs Unbiased Estimators

An estimator is said to be unbiased if its expectation value is equal to the true value of the parameter. For example, if our estimator for the mean is the arithmetic average $\hat{\mu} = \frac{1}{n} \sum_{i=1}^{n} X_i$ (sample mean), the arithmetic average is said to be an unbiased estimator of the mean if $E \hat{\mu}$ is equal to the true mean of $X$. The latter is defined as $\int x p_X(x) dx$. 

4.2.1. Random Sample. Suppose we have a random variable $X$. In the laboratory, we can measure $X$ by acquiring several measurements of $X$. Denote these measurements of $X$ by the set of values $(X_1, X_2, \ldots, X_n)$. Each $X_i$ is iid rv with the same distribution of $X$. $X_i$ are samples of $X$ measured at different points in time. We say that the set $(X_1, X_2, \ldots, X_n)$ is a random sample of $X$.

4.2.2. MLE of the Mean: Is the Estimator Biased? Let $X$ be a random variable with mean $\mu$ ($\mu = \int xp_X(x)dx$ is the true mean). Let $(X_1, X_2, \ldots, X_n)$ be a random sample of $X$. Here we show that 

$$\hat{\mu} = \frac{1}{n} \sum_{i=1}^{n} X_i,$$

is an unbiased estimator of $\mu$. $\frac{1}{n} \sum_{i=1}^{n} X_i$ is the sample mean. This statement follows from:

$$E \frac{1}{n} \sum_{i=1}^{n} X_i = \frac{1}{n} \sum_{i=1}^{n} E X_i = \frac{1}{n} \sum_{i=1}^{n} \mu = \frac{1}{n} (n \mu) = \mu.$$

To get the first equality, we used the linearity property of expectation value. In the second equality, we used the fact that $\{X_i\}$ is a random sample of $X$, whose mean is $\mu$.

4.3. Variance of the Sample Mean Estimator

The sample mean 

$$\hat{\mu} = \frac{1}{n} \sum_{i=1}^{n} X_i$$

is a rv because it is a sum of rv’s $X_1, X_2, \ldots, X_n$. It has a mean equal to the true mean of $X$, namely $\mu$, which we verified in the previous section, meaning that this particular expression for $\hat{\mu}$ is an unbiased estimator. By the LLN, it converges to the true mean as $n$ increases. We can also check its variance:

$$\text{var}(\hat{\mu}) = \frac{1}{n^2} \sum_{i=1}^{n} \text{var}(X_i) = \frac{\sigma^2}{n},$$

meaning that it becomes “narrowly distributed” as $n$ increases. This says that the estimator becomes more and more precise as $n$ increases. This is a desirable feature of estimators. If you have an estimator whose variance increases with $n$, this is not a good estimator. The square root of $\text{var}(\hat{\mu})$ represents the error in the mean; you will recognize it as the standard error (aka the standard deviation of the mean).
4.4. Bias of the Variance Estimator

Let \((X_1, X_2, \ldots, X_n)\) be a random sample of \(X\) (with mean \(\mu\) and variance \(\sigma^2\)). Here we show that

\[
\hat{\sigma}_n^2 = \frac{1}{n} \sum_{i=1}^{n} (X_i - \hat{\mu})^2,
\]

is a biased estimator of \(\sigma^2\). By definition, the variance of \(X_i\), which is also the variance of \(X\) (since \(\{X_i\}\) is a random sample of \(X\)), is:

\[
\sigma^2_{X_i} = \sigma^2_X = \mathbb{E}(X - \mu)^2 = \int (x - \mu)^2 p_X(x) \, dx.
\]

Also,\(^3\)

\[
\mathbb{E}\hat{\sigma}_n^2 = \mathbb{E}\left[\frac{1}{n} \sum_{i=1}^{n} (X_i - \hat{\mu})^2\right] = \frac{1}{n} \sum_{i=1}^{n} [(X_i - \mu) - (\hat{\mu} - \mu)]^2
\]

\[
= \frac{1}{n} \sum_{i=1}^{n} [(X_i - \mu)^2 - 2(X_i - \mu)(\hat{\mu} - \mu) + (\hat{\mu} - \mu)^2]
\]

\[
= \frac{1}{n} \left[ \left( \sum_{i=1}^{n} (X_i - \mu)^2 \right) - n(\hat{\mu} - \mu)^2 \right]
\]

\[
= \frac{1}{n} \left( \sum_{i=1}^{n} \sigma^2_i \right) - \text{var}(\hat{\mu}) = \sigma^2 - \text{var}(\hat{\mu}),
\]

and using the property,

\[
\text{var}(\sum_{i} a_i X_i) = \sum_{i} a_i^2 \text{var}(X_i),
\]

valid for independent random variables, we have:

\[
\text{var}(\hat{\mu}) = \text{var}\left(\frac{1}{n} \sum_{i=1}^{n} X_i\right) = \sum_{i=1}^{n} \frac{1}{n^2} \text{var}(X_i) = \sum_{i=1}^{n} \frac{1}{n^2} \sigma^2 = \frac{1}{n} \sigma^2.
\]

Thus,

\[
\mathbb{E}\hat{\sigma}_n^2 = \sigma^2 - \frac{1}{n} \sigma^2 = \frac{n-1}{n} \sigma^2
\]

\(^3\)To get from the second to the third line, we take the summation inside the square bracket and apply it to the second and third terms. (After multiplying the second term by \(n/n = 1\).) The third term inherits a coefficient of \(n\) because it is summed \(n\) times. In the second term, we replace \((1/n) \sum X_i\) by \(\hat{\mu}\). The second term, which becomes \(-2n(\hat{\mu} - \mu)^2\) is of the same form as the third term, \(n(\hat{\mu} - \mu)^2\). Adding the two terms together gives \(-n(\hat{\mu} - \mu)^2\). When going to the fourth line, the factor of \(n\) cancels the \(1/n\) in front of the square bracket, leaving \(-\mathbb{E}(\hat{\mu} - \mu)^2\), minus 1 times the variance of \(\hat{\mu}\).
which shows that $\hat{\sigma}_n^2$ is a biased estimator of $\sigma^2$. Thus, to get an estimator which, on the average, yields a result equal to the true value of $\sigma^2$ we should take instead $\frac{n}{n-1}\hat{\sigma}_n^2$ as the estimator of the variance. We denote this “bias-corrected” estimator of the variance by $\hat{\sigma}_{n-1}^2$:

$$\hat{\sigma}_{n-1}^2 = \frac{1}{n-1} \sum_{i=1}^{n} (X_i - \hat{\mu})^2$$

The factor $\frac{n}{n-1}$ is called Bessel’s correction.

This result is the expression for the sample variance introduced earlier in the course (see, e.g., Eq. 1.1). The sample variance is basically the MLE estimator for the variance, but corrected for bias. Finally, we note that for large $n$, $1/n$ and $1/(n-1)$ are asymptotically equal. Thus, $\hat{\sigma}_n^2$ is asymptotically unbiased, i.e.

$$\lim_{n \to \infty} \hat{\sigma}_n^2 = \hat{\sigma}_{n-1}^2.$$

4.5. MLE Can Fail

The MLE method does not always work. In fact, it can fail even in the simplest cases. Suppose that the density is a mixture of two normal densities:

$$p(x; a, \sigma) = \frac{1}{2\sigma\sqrt{2\pi}} \exp\left[-\frac{(x-a)^2}{2\sigma^2}\right] + \frac{1}{2\sqrt{2\pi}} \exp\left[-\frac{x^2}{2}\right],$$

where the parameters $(a, \sigma)$ are unknown. Suppose that we measure iidrv $X_1, \ldots, X_n$ (whose values are denoted $x_1, \ldots, x_n$). For any constant $\delta > 0$, there exists small enough $\sigma = \sigma_0$ that for $a = x_1$,

$$L(x_1, \sigma_0) = \sum_{i=1}^{n} \log p(x_i; x_1, \sigma_0)$$

$$> \log \left( \frac{1}{2\sigma_0\sqrt{2\pi}} \right) + \sum_{i=2}^{n} \log \left( \frac{1}{2\sqrt{2\pi}} \exp\left[-\frac{x_i^2}{2}\right] \right)$$

$$= - \log \sigma_0 - \left( \sum_{i=2}^{n} \frac{x_i^2}{2} \right) - n \log 2\sqrt{2\pi} > \delta$$

From this inequality, we can conclude that the likelihood does not exist (i.e. it can always be made to “blow up” by taking the limit $\sigma_0 \to 0$). Therefore, MLE is unable to estimate the parameters $a$ and $\sigma$. Thus, the range of applicability of MLE is limited.

4.6. Method of Moments

We shall say a few words about an alternative method for deriving estimators, called the method of moments. Recall the strong law (SLLN) of large
numbers (Section 2.28.1): let $X_1, \ldots, X_n$ be iidrv with finite first moment. Then,
\[
\frac{1}{n} \sum_{i=1}^{n} X_i \to \mathbb{E}(X_1) \text{ almost surely as } n \to \infty.
\]
In other words, the set of random events for which the random variable $\frac{1}{n} \sum_{i=1}^{n} X_i$ does not converge to $\mathbb{E}(X_1)$ as $n \to \infty$ has probability 0 (see the actual statement of the theorem in Section 2.28.1). A second theorem of probability theory (not proved here) says that if $Y_1, \ldots, Y_n$ is a sequence of rv’s (not necessarily iidrv), such that $Y_n \to Y$ almost surely, and if $h$ is a continuous function, then $h(Y_n) \to h(Y)$ converges almost surely as $n \to \infty$.

Let $X_1, \ldots, X_n$ be iidrv whose pdf is the exponential probability
\[
p_{\theta}(x) = \theta^{-1} e^{-x\theta^{-1}} 1_{\{x>0\}},
\]
where $1_{\{x>0\}}$ is the indicator function that equals 1 when $x > 0$ and 0 otherwise. Suppose we want to estimate $\theta = \mathbb{E}(X_1)$ from experimental data. We propose the following estimator for the first moment:
\[
\hat{\theta}_n^{(1)} = \frac{1}{n} \sum_{i=1}^{n} X_i,
\]
which is such that $\hat{\theta}_n^{(1)} \to \theta$ almost surely as $n \to \infty$. The SLLN guarantees that this estimator converges to $\mathbb{E}(X_1)$, which equals $\theta$, the desired parameter. This gives us a possible estimator for $\theta$.

On the other hand, $\mathbb{E}(X_1^2) = 2\theta^2$. According to the SLLN,
\[
\frac{1}{n} \sum_{i=1}^{n} X_i^2 \to \mathbb{E}(X_1^2) = 2\theta^2 \text{ almost surely as } n \to \infty.
\]
Thus, if we divide by 2 and take the square root, we should get another estimate for $\theta$. Let
\[
Y_n = \frac{1}{n} \sum_{i=1}^{n} X_i^2,
\]
and take $h(x) = \sqrt{x/2}$. We know that the sequence of rv’s $Y_n$ converges to the rv $Y = \mathbb{E}(X_1^2) = 2\theta^2$ almost surely per the SLLN. According to the second theorem, $h(Y_n) \to h(Y) = \theta$ almost surely as $n \to \infty$. Thus,
\[
\hat{\theta}_n^{(2)} = \sqrt{\frac{1}{2n} \sum_{i=1}^{n} X_i^2},
\]
is another possible estimator for $\theta$. So we have derived 2 estimators so far, $\hat{\theta}_n^{(1)}$ and $\hat{\theta}_n^{(2)}$; you can also derive more. Which estimator you should use
depends on your particular application. Factors that may influence your
decision include bias, errors, rate of convergence, etc.

4.7. Problems

Problem 54. Use the method of moments to find parameter estimators for
the mean and variance of a Gaussian distribution.

Problem 55. Use the method of moments to find estimate(s) of the
θ parameter for the uniform distribution, with pdf
\[ p_\theta(x) = \theta^{-1}1_{[0, \theta]}(x), \theta > 0, \]
where \( \theta \) defines the upper bound of the support of \( p_\theta(x) \), and \( E(X_1) = \theta/2 \).

Problem 56. Let \( x_1, x_2, \ldots, x_n \) \( (x_i \geq 0) \) be iidrv. The pdf for each of these
rv’s is:
\[ p(x) = \frac{\beta^\alpha}{\Gamma(\alpha)} x^{\alpha-1} e^{-\beta x}, \quad i = 1, \ldots, n \]
where \( \alpha > 0 \) and \( \beta > 0 \) are parameters for the distribution and \( \Gamma(\cdot) \) is the
gamma function. Write down the likelihood function (joint pdf) for these \( n \)
observations (the parameters \( \alpha \) and \( \beta \) are the same for all \( n \) observations,
since they are iidrv). Show that the maximum likelihood estimators for \( \alpha \)
and \( \beta \) are obtained by solving these two equations:
\[
\hat{\beta} = \frac{\hat{\alpha}}{\frac{1}{n} \sum_{i=1}^{n} x_i}, \quad \text{and} \quad \log(\hat{\alpha}) - \frac{\Gamma'(\hat{\alpha})}{\Gamma(\hat{\alpha})} = \log \left( \frac{1}{n} \sum_{i=1}^{n} x_i \right) - \frac{1}{n} \sum_{i=1}^{n} \log x_i.
\]
Explain how you would calculate the values of \( \hat{\alpha} \) and \( \hat{\beta} \) in terms of the
observations \( x_1, x_2, \ldots, x_n \).

Solution. The likelihood function is
\[ L = p(x_1, x_2, \ldots, x_n | \alpha, \beta) = \frac{\beta^n \alpha}{(\Gamma(\alpha))^n} \prod_{i=1}^{n} x_i^{\alpha-1} e^{-\beta \sum_{i=1}^{n} x_i} \]
The log likelihood function is
\[ l = \log L = n \alpha \log \beta - n \log \Gamma(\alpha) + (\alpha - 1) \left( \sum_{i=1}^{n} \log x_i \right) - \beta \sum_{i=1}^{n} x_i. \]
Differentiating this expression with respect to \( \alpha \) and \( \beta \) we get
\[
\frac{\partial l}{\partial \alpha} = n \log \beta - \frac{n}{\Gamma(\alpha)} \Gamma'(\alpha) + \sum_{i=1}^{n} \log x_i \bigg|_{\alpha, \beta = \hat{\alpha}, \hat{\beta}} = 0
\]
\[
\frac{\partial l}{\partial \beta} = \frac{n \alpha}{\beta} - \sum_{i=1}^{n} x_i \bigg|_{\alpha, \beta = \hat{\alpha}, \hat{\beta}} = 0
\]
4. Statistical Parameter Estimation

Thus,
\[ \hat{\beta} = \frac{\hat{\alpha}}{\frac{1}{n} \sum_{i=1}^{n} x_i}, \quad \text{and} \quad \log(\hat{\alpha}) - \frac{\Gamma'(-\hat{\alpha})}{\Gamma(-\hat{\alpha})} = \log \left( \frac{1}{n} \sum_{i=1}^{n} x_i \right) - \frac{1}{n} \sum_{i=1}^{n} \log x_i. \]

You need \( \alpha \) to calculate \( \beta \), but in order to calculate \( \alpha \) you have a highly nonlinear equation in \( \alpha \). There is no way you can solve for \( \alpha \) analytically.

So given values for \( x_1, \ldots, x_n \) you could use Newton Raphson (or similar method) to solve for \( \alpha \) numerically.

\[ \blacksquare \]

**Problem 57.** Suppose \( x_1, x_2, \ldots, x_n \) are iid rv with density \( p(x|\theta) = \theta e^{-\theta x}, x > 0, \theta > 0 \). Write down the likelihood function for \( n \) observations. Find the maximum likelihood estimate for \( \theta \).

**Solution.** Likelihood function is
\[ L = \theta e^{-\theta x_1} \cdot \theta e^{-\theta x_2} \cdots \theta e^{-\theta x_n} = \theta^n e^{-\theta \sum_{i=1}^{n} x_i}. \]
and using the log likelihood
\[ \log L = n \log \theta - \theta \sum_{i=1}^{n} x_i \]
differentiating with respect to \( \theta \) and setting derivative equal to zero we get
\[ \frac{n}{\theta} - \sum_{i=1}^{n} x_i = 0. \]
Thus,
\[ \hat{\theta} = \frac{n}{\sum_{i=1}^{n} x_i}. \]

\[ \blacksquare \]

**Problem 58.** Find the maximum likelihood estimate of the parameter \( \theta \) for the density \( p(x|\theta) = 1/\theta, 0 \leq x \leq \theta \), based on the measurements \( x_1, x_2, \ldots, x_n \).

**Solution.**
\[ L(\theta) = \prod_{i=1}^{n} p(x_i|\theta) = \left( \frac{1}{\theta} \right)^n. \]
This function will be maximized by choosing \( \theta \) as small as possible, subject to the restriction \( 0 \leq x_i \leq \theta, i = 1, \ldots, n \). The smallest possible value of \( \theta \) that satisfies these inequalities is clearly the largest value of the \( x_i \). Thus, the MLE of \( \theta \) is given by
\[ \hat{\theta} = \max\{x_1, x_2, \ldots, x_n\}. \]

\[ \blacksquare \]
Problem 59. Take \( x_1, x_2, \ldots, x_n \) to be iidrv each with a modified Lorentz distribution, whose pdf is:

\[
p(x|\theta) = \frac{1}{\pi} \frac{\epsilon}{(x-\theta)^2 + \epsilon^2}
\]

with unknown median \( \theta \) and known spread \( \epsilon > 0 \). Derive a maximum likelihood estimator for \( \theta \).

Solution. Differentiating the likelihood function gives

\[
\frac{\partial p(x_1, \ldots, x_n|\theta)}{\partial \theta} = 2 \frac{(x_1 - \theta)}{(x_1 - \theta)^2 + \epsilon^2} + \cdots + 2 \frac{(x_n - \theta)}{(x_n - \theta)^2 + \epsilon^2} = 0.
\]

This is a nasty equation to solve for \( \theta \). We could do it by computer by implementing an iterative process for solving the nonlinear equations. Suppose that \( \theta \) is a solution. Then the terms in the equation corresponding to data points \( x_i \) that are far from \( \theta \) are close to zero. The terms in the equation corresponding to data points \( x_i \) that are close to \( \theta \) then each have magnitude about \( (x_i - \theta)/\epsilon^2 \). So the \( \theta \) solution is, roughly speaking, a sample mean of part of the data, leaving out the more extreme values. While this estimator is more efficient than the sample median, the sample median begins to look attractive from the point of view of convenience.

Problem 60. Show that the following random variable

\[
X = e^{\mu + \sigma Z}
\]

where \( Z \) is a standard normal rv, i.e. \( Z \sim N(0, 1) \), has the pdf:

\[
p(x|\mu, \sigma) = \frac{e^{-(\log x-\mu)^2/2\sigma^2}}{\sqrt{2\pi\sigma x}}, \quad x > 0.
\]

Find the MLE for the distribution parameters \( \mu \) and \( \sigma \), given a set of \( x_1, x_2, \ldots, x_n \) measurements of the random variable \( X \). Each \( x_i \) are iidrv with the above pdf for \( X \).

Solution. We proceed as usual:

\[
P(X < x) = P(e^{\mu + \sigma Z} < x) = P(\mu + \sigma Z < \log x) = P\left(Z < \frac{\log x - \mu}{\sigma}\right)
\]

\[
= \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{\log x - \mu/\sigma} e^{-\xi^2/2}d\xi.
\]

which we recognize as being

\[
\Phi\left(\frac{\log x - \mu}{\sigma}\right)
\]
Differentiation with respect to \( x \) gives the pdf:

\[
p(x) = \frac{e^{-(\log x - \mu)^2/2\sigma^2}}{\sqrt{2\pi\sigma x}}.
\]

The likelihood function is

\[
L(\mu, \sigma) = \prod_{i=1}^{n} \frac{e^{-(\log x_i - \mu)^2/2\sigma^2}}{\sqrt{2\pi\sigma x_i}}.
\]

The log likelihood is:

\[
l(\mu, \sigma) = \sum_{i=1}^{n} \left\{ -\frac{1}{2} \log(2\pi) - \log(\sigma) - \log(x_i) - (\log x_i - \mu)^2/2\sigma^2 \right\}.
\]

Differentiation with respect to \( \sigma \) and \( \mu \) yields the following MLE:

\[
\hat{\mu} = \frac{\sum_{i=1}^{n} \log x_i}{n} \quad \text{and} \quad \hat{\sigma}^2 = \frac{\sum_{i=1}^{n} (\log x_i - \hat{\mu})^2}{n}.
\]

Problem 61. Suppose that you make \( n \) measurements of the random variable \( X \), knowing the \( X \) is Poisson distributed with Poisson parameter \( \lambda \), and \( m \) measurements of another random variable \( Y \), which is known to be exponentially distributed, i.e. its pdf is \( p(y) = a \cdot \exp(-ay) \), \( a > 0 \). Find estimators for the parameters \( \lambda \) and \( a \) in terms of the experimental data \( X_1, \ldots, X_n \) and \( Y_1, \ldots, Y_m \). (All these rv’s can be assumed independent.)

Solution. Likelihood function:

\[
L(\lambda, a) = \frac{e^{-\lambda\sum x_i} \prod x_i!}{x_1! x_2! \cdots x_n!} a^m e^{-a(\sum y_i)}
\]

Taking derivative of \( \log L \) with respect to \( \lambda \):

\[-n + (x_1 + \cdots + x_n) \frac{1}{\lambda} = 0\]

with respect to \( a \):

\[m \frac{1}{a} - (y_1 + \cdots + y_m) = 0\]

The first equation gives:

\[\hat{\lambda} = \frac{1}{n} (x_1 + \cdots + x_n)\]

The second gives:

\[\hat{a} = \frac{m}{(y_1 + \cdots + y_m)}\]

Problem 62. The maximum likelihood estimator (MLE) is the set of parameters denoted \( \hat{\theta} \) for which the likelihood function is maximized (provided
that one or more maxima exists):

\[ \hat{\theta} \in \left\{ \arg \max_{\theta} L(\theta|x_1, \ldots, x_n) \right\} \]

Alternatively, we can also maximize the log-likelihood function

\[ \hat{\theta} \in \left\{ \arg \max_{\theta} l(\theta|x_1, \ldots, x_n) \right\} \]

(a) A MLE estimate is the same regardless of whether we maximize the likelihood or the log-likelihood function. Why?

(b) In the case of a Gaussian distribution write down the log likelihood function.

(c) Carry out the maximization of the likelihood function (L) and the log-likelihood function (l) in the case of a Gaussian distribution. Do you get the same result? (Recall from (a) that you should get the same result.) Which procedure is simpler?

(d) In the case of Poisson distribution distributed random variables (the case of discrete random variables, but whose parameter(s) are continuous variables!), carry out the same MLE procedure and determine the maximum likelihood estimator for the mean and variance. For simplicity, assume that the random variables measured (let’s call them \(n_1, n_2, \ldots, n_N\)) are iidrv. This exercise teaches how to compute “sample mean” and “sample variance” when performing repeated measurements of a Poisson-distributed random variable. Does the ML prescription require you to compute an arithmetic mean?

**Solution.** (a) Because log is a monotonically increasing function. Consequently, a maximum of the likelihood function is also a maximum of the log-likelihood function.

(b) The joint Gaussian for iidrv is

\[ L(\bar{X}, \sigma^2|x_1, \ldots, x_n) = \frac{1}{(2\pi)^{n/2}\sigma^n} e^{-\chi^2/2}, \quad \text{where} \quad \chi^2 = \sum_{i=1}^{n} \frac{(x_i - \bar{X})^2}{\sigma^2} \]

taking the log we get

\[ l = -(n/2) \log 2\pi - n \log \sigma - \chi^2/2 \]

Here maximization of \(L\) corresponds to the minimization of \(\chi^2\) because of the negative sign in front of \(\chi^2\).
(c) We have done $L$ in class already. The case of $l$ gives:

$$\frac{\partial l}{\partial \sigma} = -\frac{n}{\sigma} + \frac{1}{\sigma} \chi^2 = 0$$

which yields the same result as for $L$:

$$\sigma^2 = \frac{1}{n} \sum_{i=1}^{n} (x_i - \bar{X})^2$$

And likewise, the partial with respect to $\bar{X}$ yields the arithmetic mean, as can be seen by inspection. In this case, using the log likelihood is the simpler method to use.

(d) RVs are iidrv, so the joint density is a product of individual Poisson’s. Denoting their mean as $\theta$ (in class, we used $\pi$). ML method applied to the log likelihood:

$$\frac{\partial}{\partial \theta} \left( \log \prod_{i=1}^{N} \left( e^{-\theta \theta^{n_i}} \right) \right) = \frac{\partial}{\partial \theta} \left( \sum_{i=1}^{N} [-\theta + n_i \log \theta - \log n_i!] \right) = \sum_{i=1}^{N} \left( -1 + \frac{n_i}{\theta} \right)$$

which we set = 0. This yields the arithmetic mean:

$$\theta = \frac{1}{N} \sum_{i=1}^{N} n_i. \quad \blacksquare$$

Problem 63. Let $X$ be distributed according to the distribution: $\mathbb{P}(X = 2) = \theta$, $\mathbb{P}(X = 3) = 2\theta$ and $\mathbb{P}(X = 1) = 1 - 3\theta$. (The state space is discrete: $X \in \{1, 2, 3\}$.) This distribution has a single parameter $\theta$.

(a) What is the range of allowed values for $\theta$? Find the mean and variance of $X$ (in terms of $\theta$).

(b) Suppose you want to estimate the parameter $\theta$ using $n$ random samples of $X$. Let $n_1, n_2$ and $n_3$ be the number of times $X$ equals 1, 2 or 3, respectively, within this random sample. What is the probability law for $n_1$?

(c) Find the maximum likelihood estimator $\hat{\theta}$ for $\theta$. Calculate the mean and variance of $\hat{\theta}$. Is the estimator biased?

Solution. (a) $\theta$ can take values in the interval $[0, 1/3]$, so that these probabilities must remain between 0 and 1. The mean is

$$\mathbb{E}(X) = 1 - 3\theta + 2\theta + 3 \times 2\theta = 1 + 5\theta.$$
4.7. Problems

variance is

\[ \text{var}(X) = 1 - 3\theta + 4\theta + 9 \times 2\theta - (1 + 5\theta)^2 = \theta(9 - 25\theta). \]

(b) \( n_1 \) follows a binomial law, with parameters \( n \) and \( p = 1 - 3\theta \).

(c) The likelihood function is

\[ L(\vec{X}, \theta) = \prod_{i=1}^{n} \mathbb{P}(X = x_i) = (1 - 3\theta)^{n_1} \times \theta^{n_2} \times (2\theta)^{n_3} \]

Minimizing the log of \( L \) with respect to \( \theta \),

\[ \frac{\partial}{\partial \theta} \log L = -\frac{3n_1}{1 - 3\theta} + \frac{n_2}{\theta} + \frac{2n_3}{2\theta} = -\frac{3n_1}{1 - 3\theta} + \frac{n - n_1}{\theta} = 0. \]

Thus, we conclude that the estimator is

\[ \hat{\theta} = \frac{n - n_1}{3n}. \]

Its mean is

\[ \mathbb{E}(\hat{\theta}) = \frac{1}{3} - \frac{1}{3n}(1 - 3\theta)n = \theta \]

and variance is

\[ \text{var}(\hat{\theta}) = \frac{1}{9n^2} n(1 - 3\theta)3\theta = \frac{\theta(1 - 3\theta)}{3n}. \]

Thus, \( \hat{\theta} \) is an unbiased estimator. It also converges (variance decreases with \( n \)).

Problem 64. Tesla Motors has a plant in Nevada that manufactures car batteries. If a battery has length greater than \( a \), it will not fit into the car, and this is a big problem because the cost of fixing such defective batteries is high. Tesla has been plagued recently regarding problems in its assembly line, related to poor quality control. You were tasked to visit the plan in your capacity as inspector, and do a statistical analysis of the assembly line. The first question that comes to mind is what is the probability that a battery (chosen at random in the assembly line) has length greater than \( a \), i.e. \( \mathbb{P}(X \geq a) \). (Let’s assume a 1D problem.) Show that \( a \) times this probability can never be larger than the mean of \( X \) (i.e., \( \mathbb{E}(X) \)). What is the significance of this statement?

Solution.

\[ \mathbb{E}(X) = \int_{0}^{\infty} xp(x)dx \geq \int_{a}^{\infty} xp(x)dx \geq \int_{a}^{\infty} a p(x)dx = a \mathbb{P}(X \geq a) \]

Problem 65. Suppose that you make \( n \) measurements of the random variable \( X \), knowing the \( X \) is Poisson distributed with Poisson parameter \( \lambda \),
and $m$ measurements of another random variable $Y$, which is known to be exponentially distributed, i.e. its pdf is $p(y) = a \cdot \exp(-ay)$, $a > 0$. Find estimators for the parameters $\lambda$ and $a$ in terms of the experimental data $X_1, \ldots, X_n$ and $Y_1, \ldots, Y_m$. (All these rv’s can be assumed independent.)

**Solution.** Likelihood function:

$$L(\lambda, a) = \frac{e^{-\lambda x_1 \cdots x_n}}{x_1!x_2!\ldots x_n!} a^m e^{-a(y_1 + \cdots + y_m)}$$

Taking derivative of log $L$ with respect to $\lambda$:

$$-n + (x_1 + \cdots + x_n) \frac{1}{\lambda} = 0$$

with respect to $a$:

$$m \frac{1}{a} - (y_1 + \cdots + y_m) = 0.$$

The first equation gives:

$$\hat{\lambda} = \frac{1}{n} (x_1 + \cdots + x_n).$$

The second gives:

$$\hat{a} = \frac{m}{(y_1 + \cdots + y_m)}.$$

---

**Problem 66.** What is the meaning of “bias” in the context of a statistical estimator? How does one determine the amount of bias in an estimator? What desirable properties should the estimator possess? Give an example of a biased estimator and explain what makes it biased.

**Solution.** Biased, since the coefficient is $1/n$ instead of $1/(n - 1)$.

---

**Problem 67.** The Richter magnitude of an earthquake is determined from the logarithm of the amplitude of waves recorded by seismographs (adjustments are included to compensate for the variation in the distance between the various seismographs and the epicenter of the earthquake). The formula is:

$$M_L = \log_{10}(A/A_0(\delta))$$

where $A$ is the maximum excursion of the Wood-Anderson seismograph, the empirical function $A_0$ depends only on the epicentral distance of the station, $\delta$. Both $A$ and $\delta$ are prone to measurement error. Find the uncertainty in $M_L$ due to errors in $A$ and $\delta$.

**Solution.** Propagation of error formula is:

$$\sigma_{M_L}^2 = \left| \frac{\partial M_L}{\partial A} \right|^2 \sigma_A^2 + \left| \frac{\partial M_L}{\partial \delta} \right|^2 \sigma_\delta^2,$$
where the derivatives are:

\[
\frac{\partial M_L}{\partial A} = \frac{1}{\log(10)} \frac{A_0(\delta)}{A} \frac{1}{A_0(\delta)} = \frac{1}{A \cdot \log(10)}
\]

\[
\frac{\partial M_L}{\partial \delta} = \frac{1}{\log(10)} \frac{A_0(\delta)}{A} \cdot \frac{1}{[A_0(\delta)]^2} \frac{\partial A_0(\delta)}{\partial \delta} = -\frac{\partial A_0(\delta)/\partial \delta}{A \cdot A_0(\delta) \cdot \log(10)}
\]

**Problem 68.** The fluorescence decay of a fluorophore molecule is modeled by:

\[
I(t) = \tau^{-1} e^{-t/\tau},
\]

where \(\tau\) is the fluorescence decay lifetime, or equivalently, \(\tau^{-1}\) is the decay rate. (You can think of this model as giving the probability of a molecule chosen randomly from an ensemble to be found in the excited state.) To some extent, the decay lifetime is indicative of the immediate chemical environment of the molecule. Consequently, you decide to investigate its potential use as a chemical sensor by measuring lifetime in the presence of different types of solvents and solutes. Your experiments returns values of \(t\) (i.e., \(t_1, t_2, \ldots, t_n\)) at which the fluorescence intensity drops to \(1/e\) of its initial value; i.e., the point where \(t = \tau\). Such measurements of \(t_1, t_2, \ldots, t_n\) are subject to considerable errors. To reduce variability you repeat the measurement many times and derive a suitable average.

(a) Explain how you would estimate an average lifetime in terms of the experimentally measured lifetimes \(t_1, t_2, \ldots, t_n\) which maximizes the chances of observing this experimentally measured data.

(b) If the scientific journal you are trying to publish your results in asks you to report average decay rates instead of lifetimes, how would you estimate average decay rates from the data?

**Solution.** (a) Since \(t_1, t_2, \ldots, t_n\) are iid rv, the likelihood function is a product:

\[
L(t_1, t_2, \ldots, t_n) = \tau^{-1} e^{-t_1/\tau} \ldots \tau^{-1} e^{-t_n/\tau} = \tau^{-n} \exp \left( -\tau^{-1} \sum t_i \right).
\]

The log likelihood is

\[
\log L = -n \log \tau - \tau^{-1} \sum t_i.
\]

Differentiating with respect to \(\tau\) and setting equal to zero:

\[
-\frac{n}{\tau} + \frac{1}{\tau^2} \sum t_i = 0.
\]

Solving for \(\tau\) we get the MLE which is the arithmetic average of the measured lifetimes:

\[
\hat{\tau} = \frac{1}{n} \sum t_i.
\]
(b) Decay rate is inverse of lifetime:

\[ \hat{\tau}^{-1} = \frac{n}{\sum t_i} \]

Problem 69. Suppose you record the number of cars crossing some intersection going northbound at regular intervals of 5 minutes after midnight, i.e. when there is very little traffic, so you can assume that the events of cars crossing the intersection are Poisson distributed. Let \( n_i \) the number of cars recorded in the \( i \)th time interval. Your job description requires you to produce a daily report to the city by 10:00 am which includes the data recorded along with some basic statistical analysis. In particular, the city wants you to state on the cover page the mean number of vehicles and the variance in a 5 minute time interval. However, because the numbers are Poisson distributed you are unsure if you are allowed to use the formula for sample mean, since you may recall it was derived using the assumption of Gaussian statistics — not Poisson. Derive the correct formula for the sample mean and variance in the case of Poisson statistics.

Solution. See Problem 62.

Problem 70. You measure a signal, \( y \), that is the sum of a function \( f \) contaminated by additive noise \( \xi \). In discrete form:

\[ y_i = f(x_i, \beta) + \xi_i \]

where \( \beta \) is a parameter for the function \( f \) that we wish to obtain. (The form of the function \( f \) is known, but we must determine this parameter from the data.) The density of the noise \( \xi_i \) is known, \( p(\xi) \). Estimate the function \( f(x, \beta_0) \) from the set of functions \( f(x, \beta) \) (i.e. determine the value of \( \beta \)), using the data obtained by measurements \( y \) of the function \( f \) corrupted by noise \( \xi \).

Solution. The data is given by the pairs:

\( (x_1, y_1), \ldots, (x_n, y_n) \)

We estimate \( \beta_0 \) using the ML method by maximizing the log likelihood:

\[ l(\beta) = \sum_{i=1}^{n} \log p(y_i - f(x_i, \beta)) \]

Where \( p(\xi) \) is a known function and \( \xi = y - f(x, \beta) \). To make further progress we must know the form of \( p(\xi) \). If \( p \) is Gaussian,

\[ p(\xi) = \frac{1}{\sqrt{2\pi\sigma^2}} \exp \left( -\frac{\xi^2}{2\sigma^2} \right) \]
(zero mean, known variance), we then obtain the least squares method:

\[ l(\beta) = -\frac{1}{2\sigma^2} \sum_{i=1}^{n} (y_i - f(x_i, \beta))^2 - n \log(\sqrt{2\pi\sigma}) \]

Maximizing \( l(\beta) \) over the parameters \( \beta \) is the same as minimizing the least squares functional:

\[ \sum_{i=1}^{n} (y_i - f(x_i, \beta))^2 \]

The particular value of \( \beta \) that minimizes this functional is called \( \beta_0 \). For other distributions, \( p(\xi) \), a different functional may be obtained.
Data Fitting

5.1. Linear Least Squares

An important aspect of experimental research is the measurement of physical quantities. From these physical quantities we seek to confirm or disprove certain hypotheses. This could be, for example, verification that a theory holds in a certain regime. We then need to “fit” the data to an equation and determine the unknown coefficients in the equation. This is the topic of data fitting.

5.1.1. Least Squares Method. An old and trusted method to data fitting is the method of least squares. If the distance between the experimental data and the model is normally distributed with mean 0 and finite variance, the joint pdf describing the measurement of the data points \( \{y_1, \ldots, y_n\} \) is such that the application of the principle of maximum likelihood estimation (MLE) for the model parameters yields the method of least squares. In contrast, if the measurements are not normally distributed about the model, MLE does not yield least squares. (Exercise: can you demonstrate this?) In this course, we will limit our discussion to data points which are normally distributed about the model. Linear least squares is a special case of least squares when the model is linear in the fitting parameters. Non-linear least squares problems are more complicated and generally cannot be solved analytically. Later in the course, we will solve non-linear least squares problems using computer-based methods.

5.1.2. Straight line. Let us begin with the method of linear least squares. And for simplicity we shall consider the problem of fitting data to a straight
line model. Suppose we have the data shown in Fig. 5.1 and we would like to fit these points to a straight line

\[ y(x) = A + Bx. \]

The measured data is the set of pairs

\[ (x_1, y_1), (x_2, y_2), (x_3, y_3), \ldots, (x_n, y_n). \]

If we assume that

- The uncertainty in \( \{x_i\} \) is negligible.
- Each \( y_i \) is Gaussian-distributed with the same width \( \sigma_y \).
- All measurements \( \{y_i\} \) are statistically independent.

The hope is that the model (if correct) gives the true value of \( y \):

\[ y(x_i) \equiv (\text{true value of } y_i) = A + Bx_i. \]

From the above assumptions, it follows that the probability of a single measurement \( y_i \) is

\[ p(y_i) = \frac{1}{\sqrt{2\pi\sigma_y}} \exp \left[ -\frac{(y_i - A - Bx_i)^2}{2\sigma_y^2} \right]. \]

Because of statistical independence the joint probability density, \( L(A, B) \equiv p(y_1, \ldots, y_n) \), of obtaining a complete set of measurements \( y_1, \ldots, y_n \) is the product of probabilities of individual measurements:

\[ p(y_1) \cdot p(y_2) \cdots p(y_n) = \frac{1}{(2\pi)^{n/2}\sigma_y^n} \exp \left( -\sum_{i=1}^{n} \frac{(y_i - A - Bx_i)^2}{2\sigma_y^2} \right) \]

by independence

by assumption of independent Gaussians

---

1Linear least squares and the straight line model are two different things. Fitting a polynomial is still a linear least squares problem because linear means linear in the model parameters. It is not a statement about the degree of the polynomial being fitted.
We shall rewrite this expression in slightly more convenient form:

\[
L(A, B|y_1, \ldots, y_n) \equiv p(y_1, \ldots, y_n) = \frac{1}{(2\pi)^{n/2}\sigma_y^n} e^{-\chi^2/2},
\]

where

\[
\chi^2 = \sum_{i=1}^{n} \frac{(y_i - A - Bx_i)^2}{\sigma_y^2} = \sum \tilde{R}_i^2.
\]

Here, \( \tilde{R}_i \equiv (y_i - y(x_i))/\sigma_i \sim N(0, 1) \).

According to the principle of maximum likelihood, the best estimates for the unknown constants \( A \) and \( B \) are those for which \( p(y_1, \ldots, y_n) \) is maximized. Or equivalently, for which \( \chi^2 \) is a minimum. This strategy is the basis for the method of least squares fitting.

\[
\frac{\partial \chi^2}{\partial A} = -2 \frac{\sigma_y^2}{\sigma_y^2} \sum_{i=1}^{n} (y_i - A - Bx_i) = 0
\]

\[
\frac{\partial \chi^2}{\partial B} = -2 \frac{\sigma_y^2}{\sigma_y^2} \sum_{i=1}^{n} x_i(y_i - A - Bx_i) = 0
\]

which we rewrite as

\[
A \sum x_i + B \sum x_i^2 = \sum y_i
\]

\[
A \sum x_i + B \sum x_i y_i = \sum x_i y_i
\]

These 2 equations and 2 unknowns are easily solved to yield:

\[
A = \frac{\sum x_i^2 \sum y_i - \sum x_i \sum x_i y_i}{\Delta}, \quad B = \frac{n \sum x_i y_i - \sum x_i \sum y_i}{\Delta}
\]

where

\[
\Delta = n \sum x_i^2 - (\sum x_i)^2.
\]

The resulting line \( y = A + Bx \) is called the least squares fit to the data, or the line of regression of \( y \) on \( x \).

However, \( A \) and \( B \) are derived from experimental data. Thus, there is uncertainty associated with \( A \) and \( B \). Another question of interest is how do we obtain an estimate for \( \sigma_y \) based on experimental data. \( \sigma_y \) is the parameter of the (assumed) Gaussian distribution that describes the fluctuations of the experimental data about the model.

5.1.3. Solving Systems of 2 Equations and 2 Unknowns. To solve for \( N \) unknowns in \( N \) equations we can use the matrix inverse. The formula

\footnote{The method is called least squares because it involves minimizing chi square. This is a consequence of our assumption that the data is normally distributed about the model. Because of the particular form of \( L \) obtained, maximizing \( L \) is equivalent to minimizing \( \chi^2 \).}
The inverse of a $2 \times 2$ matrix is:
\[
\begin{pmatrix} a & b \\ c & d \end{pmatrix}^{-1} = \frac{1}{ad-bc} \begin{pmatrix} d & -b \\ -c & a \end{pmatrix}.
\]

The cross product $ad-bc$ is the determinant of the $2 \times 2$ matrix. Students unfamiliar with matrix inverses should check that the inverse of this $2 \times 2$ matrix $A = \begin{pmatrix} a & b \\ c & d \end{pmatrix}$, denoted $A^{-1}$, satisfies the following conditions:

$AA^{-1} = I$ and $A^{-1}A = I$. Here, $I$ is the $2 \times 2$ identity matrix, $I = \begin{pmatrix} 1 & 0 \\ 0 & 1 \end{pmatrix}$.

In the previous section we need to solve:

An + B \sum x_i = \sum y_i
A \sum x_i + B \sum x_i^2 = \sum x_i y_i.

This is done by rewriting it in matrix form:

\[
\begin{pmatrix} A \\ B \end{pmatrix} = \left( \begin{pmatrix} n \\ \sum x_i \sum x_i^2 \end{pmatrix} \right)^{-1} \left( \begin{pmatrix} \sum y_i \\ \sum x_i y_i \end{pmatrix} \right).
\]

Multiplying both sides on the left by the inverse of \( \left( \begin{pmatrix} n \\ \sum x_i \sum x_i^2 \end{pmatrix} \right) \), we solve for $A, B$:

\[\begin{pmatrix} A \\ B \end{pmatrix} = \left( \begin{pmatrix} n \\ \sum x_i \sum x_i^2 \end{pmatrix} \right)^{-1} \left( \begin{pmatrix} \sum y_i \\ \sum x_i y_i \end{pmatrix} \right), \tag{5.1}\]

Here, the inverse of that matrix is:

\[\left( \begin{pmatrix} n \\ \sum x_i \sum x_i^2 \end{pmatrix} \right)^{-1} = \frac{1}{\Delta} \left( \begin{pmatrix} \sum x_i^2 \\ -\sum x_i \\ n \end{pmatrix} \right). \tag{5.2}\]

Substitution of (5.2) into (5.1) gives the final result:

\[
\begin{pmatrix} A \\ B \end{pmatrix} = \frac{1}{\Delta} \left( \begin{pmatrix} \sum x_i^2 \sum y_i - \sum x_i \sum x_i y_i \\ n \sum x_i y_i - \sum x_i \sum y_i \end{pmatrix} \right), \quad \Delta = n \sum x_i^2 - (\sum x_i)^2.
\]

### 5.1.4. Estimating $\sigma_y$ for a Straight Line Model.

The deviations $y_i - A - Bx_i$ are normally distributed, all with mean 0 and width $\sigma_y$. We can derive an expression for $\sigma_y$ by viewing $L$ as a function of $\sigma_y$ and use maximum likelihood. Thus, we set:

\[
\frac{\partial}{\partial \sigma_y} \left( \frac{1}{\sigma_y^n} e^{-\chi^2/2} \right) = 0, \quad \text{where} \quad \chi^2 = \sum_{i=1}^{n} \frac{(y_i - A - Bx_i)^2}{\sigma_y^2}.
\]
Carrying out the differentiation we get
\[ -n \sigma_y e^{-\chi^2/2} + \frac{1}{\sigma_y} \left( -\frac{1}{2} \right) \left( -\frac{2}{\sigma_y^2} \right) \sum_{i=1}^{n} (y_i - A - B x_i)^2 = 0 \]
from which we conclude that \( \sigma_y^2 = \frac{1}{n} \sum_{i=1}^{n} (y_i - A - B x_i)^2 \). In practice we take the prefactor to be \( 1/(n-2) \) instead of \( 1/n \) because \( A \) and \( B \) are computed from the data and this results in a bias of the estimator \( \sigma_y^2 \) (verify this!). A prefactor of \( 1/(n-2) \) corrects this bias. This “fix” makes sense if we imagine fitting a straight line to only \( n=2 \) data points. The fit would always be perfect (\( \sigma_y = 0 \)). The prefactor \( 1/(n-2) \) ensures that we get a division by zero: \( \sigma_y = 0/0 \) (undefined), indicating that such a situation should be avoided.

\[ \sigma_{y, n-2}^2 = \frac{1}{n-2} \sum_{i=1}^{n} (y_i - A - B x_i)^2. \]

### 5.1.5. Estimating the Magnitude of \( \sigma_A \) and \( \sigma_B \)

For simplicity let us assume that the errors in \( \{y_i\} \) are all identical, i.e. \( \sigma_{y_i} = \sigma_y \). Here we cannot use maximum likelihood because \( L(A, B, \sigma_y) \) does not depend on \( \sigma_A \) or \( \sigma_B \). However, since \( A \) and \( B \) are well-defined functions of \( y_1, y_2, \ldots, y_n \), we can find \( \sigma_A \) and \( \sigma_B \) by simple error propagation:

\[
A = \frac{\sum x_i^2 \sum y_i - \sum x_i \sum x_i y_i}{\Delta}, \quad \Delta = n \sum x_i^2 - (\sum x_i)^2
\]

\[
\sigma_A^2 = \sum_{j=1}^{n} \left| \frac{\partial A}{\partial y_j} \right|^2 \sigma_{y_j}^2 = \sum_{j} \left| \frac{\sum x_i^2 - x_j \sum x_i}{\Delta} \right|^2 \sigma_y^2
\]

\[
= \sum_{j} \left[ (\sum x_i^2)^2 - 2 \sum x_i^2 x_i^2 + x_i^2 \sum x_i^2 \right] \sigma_y^2 / \Delta^2
\]

\[
= \left[ n (\sum x_i^2)^2 - 2 \sum x_i^2 (\sum x_i)^2 + (\sum x_i^2)(\sum x_i)^2 \right] \sigma_y^2 / \Delta^2
\]

\[
= \left[ \sum x_i^2 \left[ n \sum x_i^2 - (\sum x_i)^2 \right] \right] \sigma_y^2 / \Delta^2
\]

from which we get

\[ \sigma_A = \sigma_y \sqrt{\frac{\sum x_i^2}{\Delta}}. \]

\^It is worth reminding ourselves of our assumption that the error in \( \{x_i\} \) is negligible. If this error is not negligible, we need to modify our strategy.
We proceed the same manner for $\sigma_B$.

Given:

$$ B = \frac{n \sum x_i y_i - \sum x_i \sum y_i}{\Delta}, \quad \Delta = n \sum x_i^2 - (\sum x_i)^2 $$

we apply the method of error propagation to derive the error in $B$:

$$ \sigma_B^2 = \sum_j \left| \frac{\partial B}{\partial y_j} \right|^2 \sigma_y^2 = \sum_j \left| \frac{n x_j - \sum x_i}{\Delta} \right|^2 \sigma_y^2 $$

$$ = \sigma_y^2 \sum_j \frac{n^2 x_j^2 - 2 n x_j \sum x_i + (\sum x_i)^2}{\Delta^2} $$

$$ = n \left[ \Delta \sum_j x_j^2 - (\sum x_i)^2 \right] \frac{\sigma_y^2}{\Delta^2} = \frac{n \sigma_y^2}{\Delta} $$

from which we conclude that

$$ \sigma_B = \sigma_y \sqrt{\frac{n}{\Delta}}. $$

5.1.6. Factors that Affect a Straight Line Fit. Suppose that we fit data to the straight line:

$$ y(x_i; A, B) = A + Bx_i $$

A number of factors will affect the results (slope and intercept).

- **Intercept.** To increase the precision in the intercept, we must reduce the error bars on points that are close to the $y$-axis.

- **Slope.** To increase the precision of the slope, we must reduce the error bars on the points located at extrema of the data set (first and last points).

This is illustrated in Figure 5.2.
5.2. How to Determine if a Fit is Good: Inspect the Residuals for Possible Trends

The residuals measure the distance between the model, \( \{y(x_i)\} \), and the data, \( \{y_i\} \):

\[
R_i = y_i - y(x_i; \tilde{\beta})
\]

where \( \tilde{\beta} \) denotes the set of fitting parameters (A, B in the case of our straight line model). Plotting the residuals (see Fig. 5.4) yields a useful tool to judge if a fitting model is adequate. In the figure below we have illustrated two examples where the fitting model is inadequate because the residuals show some clear (non-random) trends. In the first case, a quadratic dependence should be added to the model. In the second case, the model fails to capture a sinusoidal dependence.

Suppose that \( y(x) \) is a model for the experimental data. The set of ordered pairs, \( \{(x_i, y_i)\} \), is the experimental data. For simplicity, we still assume that the errors in \( x_i \) are negligible. So far we have assumed that the following random variable is normally distributed about the model’s trendline (with parameters: mean, \( y(x_i) \) and variance, \( \sigma^2_y \)):

\[
y_i \sim N(y(x_i), \sigma^2_y).
\]

An equivalent statement is:

\[
R_i = y_i - y(x_i) \sim N(0, \sigma^2_y).
\]

Invoking \( \text{var}(aX) = a^2 \cdot \text{var}(X) \), another equivalent statement is:

\[
\tilde{R}_i = \frac{y_i - y(x_i)}{\sigma_y} \sim N(0, 1).
\]

where \( \sigma_y \) is the error in \( y_i \). The basic assumption of the least squares method that data points are Gaussian-distributed about the model’s trendline is a good assumption in most cases. This is a consequence of the central
limit theorem and the fact that most physical measurements of macroscopic properties are the result (sum) of a very large quantity of smaller microscopic processes, such as molecular collisions and other scattering events. There are exceptions to this, where in some cases the distribution is not Gaussian.

\( \tilde{R}_i \) is known as the normalized residuals. In the sketch below, the plot on the left shows hypothetical experimental data whose error bars are larger on the right than on the left. This leads to the residuals \( \{R_i\} \) shown in the middle plot. The error grows from left to right, as seen in the residuals. However, if we normalize the residuals, \( \{\tilde{R}_i\} \), the results show a uniform error bar since \( \tilde{R}_i \sim N(0,1) \). This is shown on the plot to the right, where I have sketched a bell-shaped histogram to illustrate the standard Gaussian distribution of the normalized residuals (Fig. 5.4).

An adequate model which captures all the needed trends should leave normalized residuals that are normally distributed with mean 0 and width 1. A good fit will yield:

- 68% of data scattered within \( \pm 1 \) of 0.
- 95% of data scattered within \( \pm 2 \) of 0.
- 99.7% of data scattered within \( \pm 3 \) of 0.

### 5.3. Chi-square as a Goodness-of-Fit Parameter

For convenience, we shall denote the parameters of the model \((A, B, C, \ldots)\) by the vector \( \beta = (A, B, C, \ldots) \). Consider a set of data points \( \{(x_i, y_i)\} \) and a model \( y(x_i; \beta) \) for the data (Fig. 5.5).

A measure of how good the fit is should be a distance metric that measures how far the data points lie from the curve. One possible such measure can be obtained by summing all differences (squared) between the data points...
and the model. The chi-square
\[ \chi^2 = \sum_{i=1}^{n} \frac{(y_i - y(x_i; \vec{\beta}))^2}{\sigma_{y_i}^2} \]
is a measure of the distance between the model and the data, i.e. the “goodness-of-fit”. Some remarks:

- Fitting the data corresponds to minimizing \( \chi^2 \). Doing so ensures that the model closely represents the measured data. The best fit parameters are found when \( \chi^2 \) is minimized.

- \( \chi^2 \) can also be viewed as a random variable since it is a function of \( y_i \) which are themselves random variables. Thus, it has its own distribution. (Exercise: can you derive its probability distribution?)

- It can lead to analytical expressions for the coefficients in some cases. For example, we have seen that for linear models, analytical expressions can be obtained.

- When analytical expressions are not possible, we can always use computer-based minimization. (Computer-based optimization is the topic of subsequent lectures.)

### 5.3.1. Chi-Square is \( (l_2\text{-norm})^2 \) of the Normalized Residuals.

The formula for chi-square
\[ \chi^2 = \sum_{i=1}^{n} \left[ \frac{y_i - y(x_i; \vec{\beta})}{\sigma_{y_i}} \right]^2 = \sum_{i=1}^{n} \tilde{R}_i^2 \]
where
\[ \tilde{R}_i = \frac{y_i - y(x_i; \vec{\beta})}{\sigma_{y_i}} \]
is the normalized residual, can be viewed also known as the \( l_2 \)-norm of a \( n \)-dimension vector whose components are the normalized residuals:

\[
\vec{R} = (\bar{R}_1, \bar{R}_2, \ldots, \bar{R}_n).
\]

The \( l_2 \) norm of a vector \( \vec{x} \) is also known as the Euclidean norm:

\[
\| \vec{x} \|_2 = \sqrt{x_1^2 + \cdots + x_n^2}.
\]

where \( \vec{x} = (x_1, x_2, \ldots, x_n) \). The Euclidean distance between two vectors \( \vec{x} \) and \( \vec{y} \) is,

\[
\| \vec{x} - \vec{y} \| = \sqrt{(x_1 - y_1)^2 + (x_2 - y_2)^2 + \cdots + (x_n - y_n)^2}.
\]

The \( l_1 \) norm of a vector \( \vec{x} \) is defined as:

\[
\| \vec{x} \|_1 = |x_1| + |x_2| + \cdots + |x_n|.
\]

The \( l_1 \) norm is also known as the Manhattan metric or the taxicab norm. The taxicab distance is thus:

\[
\| \vec{x} - \vec{y} \|_1 = |x_1 - y_1| + |x_2 - y_2| + \cdots + |x_n - y_n|.
\]

In Fig. 5.6, the taxicab distance in 2D \( (n = 2) \) is shown by the red line and the Euclidean distance is shown by the green curve.

Generally, the \( l_p \)-norm of a vector \( \vec{x} \) is:

\[
\| \vec{x} \|_p = (|x_1|^p + \cdots + |x_n|^p)^{1/p}.
\]

As you can see, there are many possible choices for the distance metric in order to measure the distance between the data and the model. With the chi-square function, as constructed, we have a \( n \)-dimensional space and least squares fitting consists of minimizing \( \chi^2 \), the \( l_2 \)-norm between the data points \( \{y_i\} \) and the model \( \{y(x_i; \beta)\} \). The \( l_2 \) norm is easier to work with from a mathematical standpoint than other distance metrics. For example, it is generally difficult to work with absolute values.
5.4. R-Squared Value: The Coefficient of Determination

The coefficient of determination, $R^2$, is defined as:

$$R^2 = 1 - \frac{SS_{res}}{SS_{tot}} = 1 - \frac{\text{unexplained variance}}{\text{total variance}} = 1 - \frac{var(\epsilon)}{var(y_i)}$$

where

$$SS_{res} = \sum_{i=1}^{n} (y_i - y(x_i))^2, \quad var(\epsilon) = \frac{1}{n} \sum_{i=1}^{n} (y_i - y(x_i))^2,$$

$$SS_{tot} = \sum_{i=1}^{n} (y_i - \bar{y})^2, \quad \bar{y} = \frac{1}{n} \sum_{i=1}^{n} y_i,$$

$\{y_i\}$ are the observed data, $\{y(x_i)\}$ are the model predictions. $R^2$ is a method that can tell us how well observed outcomes are replicated by a model. Like $\chi^2$, it also uses the sum of square differences ($L^2$ distance) between the data and the model. The better the fit, the closer the coefficient of determination gets to $R^2 = 1$.

5.4.1. Linear Least Squares: Summary of Assumptions. It is important to remember the assumptions behind the method of linear least squares. First of all, we assumed Gaussian-distributed errors. By error we mean deviations of the data from the model, $y_i - y(x_i)$. This is reflected in the use of Gaussian PDFs that describe the probability of a measurement $y_i$ being found near the model $y(x_i)$:

$$P(y_i \leq Y_i \leq y_i + dy_i) = p_{Y_i}(y_i)dy_i = \frac{1}{\sqrt{2\pi}\sigma_{y_i}^2}e^{-\frac{(y_i - y(x_i))^2}{2\sigma_{y_i}^2}}dy_i,$$

i.e., each $Y_i$ obeys a normal law with mean $y(x_i)$ and variance $\sigma_{y_i}^2$, or $Y_i \sim N(y(x_i), \sigma_{y_i}^2)$. We have also assumed that the measured data points $\{y_1, y_2, \ldots, y_n\}$ are statistically independent of each other. This assumption of statistical independence allowed us to write the likelihood function as a product of Gaussians:

$$p_{Y_1}(y_1) \cdot p_{Y_2}(y_2) \cdots \cdot p_{Y_n}(y_n) = \frac{1}{(2\pi)^{n/2} \prod_{i=1}^{n} \sigma_{y_i}} e^{-\chi^2/2}, \quad \chi^2 = \sum_{i=1}^{n} \tilde{R}_i^2.$$

Finally, the last assumption is our choice of the method of maximum likelihood estimation (MLE) as a way to derive equations for the fitting parameter. MLE involves choosing the fitting parameters that lead to the maximum likelihood of observing the experimental data $\{(x_i, y_i)\}$. 
5.5. $\chi^2$ Distribution

We note that $\chi^2$ is a function of several independent random variables:

$$\chi^2 = \sum_{i=1}^{n} \left( \frac{y_i - y(x_i)}{\sigma_i} \right)^2 = \sum_{i=1}^{n} R_i^2,$$

where $R_i$ are each normally distributed with mean 0 and variance 1. Let’s derive the distribution of $\chi^2$. In the general case of $n$ r.v’s: let $Y = X_1^2 + X_2^2 + \cdots + X_n^2$. Then,

$$P(Y < y) = P(X_1^2 + \cdots + X_n^2 < y) = \int_{\{x_1^2 + \cdots + x_n^2 < y\}} e^{-(x_1^2 + \cdots + x_n^2)/2} \frac{1}{(2\pi)^{n/2}} dx_1 \cdots dx_n.$$

The PDF is obtained by taking the limit $dy \to 0$:

$$P(y < Y < y + dy) = P(y < X_1^2 + \cdots + X_n^2 < y + dy)$$

$$= \frac{1}{(2\pi)^{n/2}} \int_{\{y < x_1^2 + \cdots + x_n^2 < y + dy\}} e^{-(x_1^2 + \cdots + x_n^2)/2} dx_1 \cdots dx_n$$

$$= \frac{y^{n/2-1}e^{-y/2}dy}{2^{n/2}\Gamma(n/2)}.$$

where $R = \sqrt{y}$, $dR = \frac{dy}{2\sqrt{y}}$ and $A = \frac{2R^{n-1} \pi^{n/2}}{\Gamma(n/2)}$ is the area of the $(n-1)$-sphere. (Recall that $\Gamma(z) = \int_0^\infty x^{z-1}e^{-x}dx$ and $\Gamma(n/2) = (n/2 - 1)!$ when $n \geq 1$ is even.)

5.5.1. Maximizing the Likelihood Function. The MLE method requires us to maximize the likelihood function $L(\hat{\beta}|y_1, y_2, \ldots, y_n)$ with respect to the fitting parameters $\beta = (A, B, C, \ldots)$. There are at least 3 ways to do this:

- Maximize $L$ by inspection. It is sometimes possible to find the maximum by inspection. For example: maximize functions of the form $(1/2 - \theta)^2$, $(1/2 + \theta)$ and $(1/2 - \theta)(1/2 + \theta)$ with respect to $\theta$, for $-1/2 \leq \theta \leq 1/2$.

---

4Recall that $P(y < Y < y + dy) = dP(Y < y) = P(Y < y + dy) - P(Y < y) \approx \frac{dy}{P(Y < y)} \cdot dy$.

5When we discuss the $\chi^2$ test later, our notation will be slightly different. We shall denote the random variable as $\chi^2$. Instead of $n$, the parameter of the distribution will be denoted as $\nu$ and $\nu$ will denote the degrees of freedom (of the $\chi^2$ distribution). Thus the random variable will be denoted as $\chi^2$. The sampling distribution of $\chi^2$ will be denoted:

$$P(\chi^2 \leq \chi^2 < \chi^2 + d(\chi^2)) = p_\nu(\chi^2)d(\chi^2) = \frac{e^{-\chi^2/2}(\chi^2)^{\nu/2-1}}{2^{\nu/2}\Gamma(\nu/2)}d(\chi^2).$$
This can be done by inspection: in the first case, \( \theta = -1/2 \); in the second case, \( \theta = 1/2 \) and in the third case, \( \theta = 0 \).

- You can maximize \( L \) using calculus. In one variable the necessary condition for an extremum is \( dL/dA = 0 \) whereas the second derivative test for a maximum is \( d^2L/dA^2 < 0 \). For example, in the case of two variables \( \beta = (A, B) \), the necessary condition for an extremum is \( L \equiv L(A, B) \) is \( dL = (\partial_A L)dA + (\partial_B L)dB = 0 \), which implies that \( \partial_A L = \partial_B L = 0 \). This gives 2 equations and two unknowns, allowing us to solve for \( A \) and \( B \).

Checking for a maximum requires a second (or higher-order) derivative test.

- You can maximize the log of \( L \), \( l = \log(L) \), also known as the “log-likelihood” function. Since log is a monotonic function, maximizing \( L \) is the same as maximizing \( \log(L) \). The advantage of working with the log of \( L \) is because log converts products into sums, i.e. \( \log(AB) = \log(A) + \log(B) \).

5.5.2. Linear Least Squares: Geometric Interpretation of \( A, B \). For the straight line model,

\[
y(x_i; A, B) = A + Bx_i,
\]

we obtained the result for the linear least squares method as:

\[
A = \frac{\sum_{i=1}^{n} x_i \sum_{i=1}^{n} y_i - \sum_{i=1}^{n} x_i \sum_{i=1}^{n} x_i y_i}{\Delta},
\]

\[
B = \frac{n \sum_{i=1}^{n} x_i y_i - \sum_{i=1}^{n} x_i \sum_{i=1}^{n} y_i}{\Delta},
\]

\[
\Delta = n \sum_{i=1}^{n} x_i^2 - (\sum_{i=1}^{n} x_i)^2.
\]

This result is rather difficult to memorize. I will show you how you can remember it using a simple geometric argument. We note that \( A \) and \( B \) can be expressed in terms of sample averages. In the equation for \( A \), we divide the numerator and denominator by \( n^2 \), and write:

\[
A = \frac{\langle x^2 \rangle \langle y \rangle - \langle x \rangle \langle xy \rangle}{\langle x^2 \rangle - \langle x \rangle^2},
\]

where the angle brackets denote sample means, i.e., \( \langle x \rangle = \frac{1}{n} \sum_{i=1}^{n} x_i \), \( \langle x^2 \rangle = \frac{1}{n} \sum_{i=1}^{n} x_i^2 \) and \( \langle xy \rangle = \frac{1}{n} \sum_{i=1}^{n} x_i y_i \). We note that \( \langle xy \rangle \) is simply an inner product of two vectors, \( \vec{x} = (x_1, \ldots, x_n) \) and \( \vec{y} = (y_1, \ldots, y_n) \), sometimes written as \( \langle \vec{x}, \vec{y} \rangle \) in a linear algebra course.
In the expression for $B$, if we divide the numerator and denominator by $n^2$, we get:

$$B = \frac{\langle xy \rangle - \langle x \rangle \langle y \rangle}{\langle x^2 \rangle - \langle x \rangle^2} = \frac{\text{cov}(x,y)}{\text{var}(x)}.$$ 

The factor $\frac{\text{cov}(x,y)}{\text{var}(x)}$ represents an orthogonal projection of the vector $\vec{y}$ onto $\vec{x}$. In linear algebra such orthogonal projections are accomplished with the use of projection operators:

$$P_{\vec{x}}(\vec{y}) = \frac{\langle \vec{y}, \vec{x} \rangle}{\|\vec{x}\|^2} \vec{x} = \langle \vec{y}, \hat{e}_x \rangle \hat{e}_x = \|\vec{y}\| \cos \theta \hat{e}_x, \quad \hat{e}_x = \frac{\vec{x}}{\|\vec{x}\|},$$

where $\theta$ is the angle between the vectors $\vec{y}$ and $\vec{x}$.

Furthermore,

$$\langle y \rangle - B \langle x \rangle = \langle y \rangle \left[ \frac{(x^2) - \langle x \rangle^2}{\langle x^2 \rangle - \langle x \rangle^2} \right] - \langle x \rangle \left[ \frac{\langle xy \rangle - \langle x \rangle \langle y \rangle}{\langle x^2 \rangle - \langle x \rangle^2} \right] = A.$$

Thus, the intercept is:

$$A = \langle y \rangle - B \langle x \rangle = \langle y \rangle - \langle x \rangle \frac{\text{cov}(x,y)}{\text{var}(x)}.$$ 

Now let’s look at our model again, $y = A + Bx$, and view $x$ and $y$ as random variables and $A$ and $B$ as constants. Taking its average, $\langle y \rangle = A + B \langle x \rangle$, yields the relationship between $A$ and $B$, $A = \langle y \rangle - B \langle x \rangle$. The equation $B = \frac{\text{cov}(x,y)}{\text{var}(x)}$ is obtained from $y = A + Bx$ by simply “projecting” it onto $x$ using the projection operator $P_{\vec{x}}(\cdot) = \frac{\text{cov}(x,\cdot)}{\text{var}(x)}$, i.e.

$$P_{\vec{x}}(\vec{y}) \equiv \frac{\text{cov}(x,y)}{\text{var}(x)} = \frac{\text{cov}(x,A + Bx)}{\text{var}(x)} = B \frac{\text{cov}(x,x)}{\text{var}(x)} = B.$$ 

### 5.6. Weighted Average

Consider the experiment where we measure a random variable $X$ using different methods. For example, $X$ could be the weight of an object and the weight can be measured using different types of balances, where each balance has its own uncertainty. Thus, we have the measurements $x_1, x_2, \ldots, x_n$, each with uncertainty $\sigma_1, \sigma_2, \ldots, \sigma_n$. The question we want to answer is what value should we report for $X$, in terms of $x_1, \ldots, x_n$ given that the uncertainties are different. Naturally, you may expect that the measurements with smaller uncertainty should carry more weight. It does not make sense to use the sample mean where all readouts are weighted equally, since some of those readouts may carry very large error bars.
5.6. Weighted Average

Let us assume that $x_1, \ldots, x_n$ are Gaussian-distributed iidrv with the same mean ($\mu$) but different variances $\sigma_i^2$. We will then find the ML estimator for the mean, $\hat{\mu}$, and an estimator for the variance of the mean, i.e. $\sigma^2_{\hat{\mu}}$, where $\mu = \hat{\mu}$. In other words, assume that the $\sigma_i^2$ are given to you (known values). The problem consists of finding $\sigma^2_{\hat{\mu}}$ in terms of the known values $x_i$ and $\sigma_i^2$.

The likelihood function is:

$$L(\mu, \sigma | x_1, \ldots, x_n) = \prod_{i=1}^{n} \frac{1}{\sqrt{2\pi \sigma_i^2}} e^{-(x_i-\mu)^2/2\sigma_i^2}.$$ 

The log-likelihood is:

$$\log L = -\frac{1}{2} n \log(2\pi) - \sum_{i=1}^{n} \log \sigma_i - \sum_{i=1}^{n} \frac{(x_i - \mu)^2}{2\sigma_i^2}.$$ 

Extrema are found from:

$$\frac{\partial (\log L)}{\partial \mu} = \sum_{i=1}^{n} \frac{(x_i - \mu)}{\sigma_i^2} = \sum_{i=1}^{n} x_i \sigma_i^{-2} - \mu \sum_{i=1}^{n} \frac{1}{\sigma_i^2} = 0.$$ 

This gives:

$$\hat{\mu} = \frac{\sum_{i=1}^{n} x_i}{\sum_{i=1}^{n} 1/\sigma_i^2}.$$ 

“weighted mean”

Note: if $\sigma_i$ are all identical (i.e. $\sigma_i = \sigma$ for all $i$) then this reduces to the simple arithmetic average, $\hat{\mu} = \frac{1}{n} \sum_{i=1}^{n} x_i$. We can get a partial check that this is a maximum of $L$:

$$\frac{\partial^2 (\log L)}{\partial \mu^2} = -\sum_{i=1}^{n} \frac{1}{\sigma_i^2} < 0.$$ 

(The “proof” that this is a maximum also requires additional considerations.)

We can check for bias:

$$E\hat{\mu} = E\frac{\sum_{i=1}^{n} x_i}{\sum_{i=1}^{n} 1/\sigma_i^2} = \sum_{i=1}^{n} \frac{E x_i}{\sigma_i^2} = \frac{\sum_{i=1}^{n} x_i}{\sum_{i=1}^{n} 1/\sigma_i^2} = \frac{\sum_{i=1}^{n} \mu}{\sum_{i=1}^{n} 1/\sigma_i^2} = \mu. \quad \text{(unbiased)}$$

The variance of the mean is obtained by error propagation

$$\sigma^2_{\hat{\mu}} = \sum_{i=1}^{n} \sigma_i^{-2} \left( \frac{\partial \mu}{\partial x_i} \right)^2.$$ 

---

6 The assumption of Gaussian distribution is made in order to simplify the math.
7 Notice how $L$ is a product of functions, whereas $l = \log L$ is a summation. Thus, working with the log-likelihood converts products into sums, which are easier to handle.
where
\[
\frac{\partial \hat{\mu}}{\partial x_i} = \frac{\partial}{\partial x_i} \sum_{i=1}^{n} \left( \frac{x_i}{\sigma_i^2} \right) = \frac{1}{\sigma_i^2} \sum_{i=1}^{n} \left( \frac{1}{\sigma_i^2} \right) ,
\]
so
\[
\sigma_{\hat{\mu}}^2 = \sum_{i=1}^{n} \sigma_i^2 \left( \frac{1}{\sum_{j=1}^{n} \left( \frac{1}{\sigma_j^2} \right) } \right)^2 = \frac{1}{\sum_{j=1}^{n} \left( \frac{1}{\sigma_j^2} \right) } = \frac{1}{\sum_{j=1}^{n} \left( \frac{1}{\sigma_j^2} \right) } .
\]
Thus, \( \sigma_{\hat{\mu}}^2 \) is equal to the harmonic mean\(^8\) of the variances of each measure-ment, \( \sigma_j^2 \), divided by \( n \):
\[
\sigma_{\hat{\mu}}^2 = \frac{1}{n} \left( \frac{\sum_{j=1}^{n} \left( \frac{1}{\sigma_j^2} \right) }{n} \right)^{-1} .
\]
Note: if \( \sigma_i \) are all identical (i.e. \( \sigma_i = \sigma \) for all \( i \)) then this reduces to
\( \sigma_{\hat{\mu}}^2 = \sigma^2 / n \), the standard error (or standard deviation of the mean).

5.7. Weighted Least Squares

If the measured data \( \{ y_i \} \) have different uncertainties \( \{ \sigma_y \} \) then we need to account for their relative “weights” when fitting the curve (Fig. 5.7), i.e. points with exceedingly large error bars should not play a dominant role in the fitting results.

We define the “weight” of the \( i \)-th measurement as \( w_i = \frac{1}{\sigma_{y_i}^2} \). We can apply the principle of maximum likelihood. We first write down the formula for

---

\(^8\)The harmonic mean \( H(x_1, \ldots, x_n) \) of \( x_1, \ldots, x_n \) is:
\[
\frac{1}{H(x_1, \ldots, x_n)} = \frac{1}{n} \left( \sum_{i=1}^{n} \frac{1}{x_i} \right) .
\]
5.7. Weighted Least Squares

chi-square:
\[ \chi^2 = \sum_{i=1}^{n} \frac{(y_i - A - Bx_i)^2}{\sigma_{y_i}^2}. \]

Differentiating,
\[
\frac{\partial \chi^2}{\partial A} = -2 \sum_{i=1}^{n} \sigma_{y_i}^{-2} (y_i - A - Bx_i) = 0
\]
\[
\frac{\partial \chi^2}{\partial B} = -2 \sum_{i=1}^{n} \sigma_{y_i}^{-2} x_i (y_i - A - Bx_i) = 0,
\]
which we can rewrite as:
\[
A \sum_{i=1}^{n} w_i + B \sum_{i=1}^{n} w_i x_i - \sum_{i=1}^{n} w_i y_i = 0
\]
\[
A \sum_{i=1}^{n} w_i x_i + B \sum_{i=1}^{n} w_i x_i^2 - \sum_{i=1}^{n} w_i x_i y_i = 0.
\]
These 2 equations in 2 unknowns can be solved to yield (see next section):
\[
A = \frac{\left( \sum_{i=1}^{n} w_i x_i^2 \right) \left( \sum_{i=1}^{n} w_i y_i \right) - \left( \sum_{i=1}^{n} w_i x_i \right) \left( \sum_{i=1}^{n} w_i y_i \right)}{\Delta}
\]
\[
B = \frac{\left( \sum_{i=1}^{n} w_i \right) \left( \sum_{i=1}^{n} w_i x_i y_i \right) - \left( \sum_{i=1}^{n} w_i x_i \right) \left( \sum_{i=1}^{n} w_i y_i \right)}{\Delta}
\]
where
\[
\Delta = \left( \sum_{i=1}^{n} w_i \right)^2 \sum_{i=1}^{n} w_i x_i^2 - \left( \sum_{i=1}^{n} w_i x_i \right)^2.
\]
We can also easily show that: \(^9\)
\[
\sigma_A = \sqrt{\frac{\sum_{i=1}^{n} w_i x_i^2}{\Delta}} \quad \text{and} \quad \sigma_B = \sqrt{\frac{\sum_{i=1}^{n} w_i}{\Delta}}.
\]

5.7.1. Solving for \( A, B \) using Matrix Inverse. For the system of equations (5.3) in the previous section, we would first write it in matrix form:
\[
\begin{pmatrix}
\sum_{i=1}^{n} w_i \\
\sum_{i=1}^{n} w_i x_i \\
\sum_{i=1}^{n} w_i x_i^2
\end{pmatrix}
\begin{pmatrix}
A \\
B \\
\sum_{i=1}^{n} w_i x_i y_i
\end{pmatrix}
= \begin{pmatrix}
\sum_{i=1}^{n} w_i y_i \\
\sum_{i=1}^{n} w_i x_i y_i
\end{pmatrix},
\]
then solve for \((AB)\) with the matrix inverse:
\[
\begin{pmatrix}
A \\
B
\end{pmatrix}
= \begin{pmatrix}
\sum_{i=1}^{n} w_i \\
\sum_{i=1}^{n} w_i x_i \\
\sum_{i=1}^{n} w_i x_i^2
\end{pmatrix}^{-1}
\begin{pmatrix}
\sum_{i=1}^{n} w_i y_i \\
\sum_{i=1}^{n} w_i x_i y_i
\end{pmatrix}.
\]

\(^9\)If you had difficulty with this type of calculation when we previously dealt with the case of identical error bars, here is a good opportunity to gain some more practice with error propagation.
The matrix inverse is easily found using the above formula:
\[
\left( \frac{1}{\sum w_i x_i^2 - (\sum w_i x_i)^2} \right) \sum w_i x_i^2 = \sum w_i (x_i - \bar{x})^2
\]
Denoting the determinant by \( \Delta \) and carrying out the multiplication of the inverse with the column vector \( \sum w_i y_i \), as required by Equation (5.4), we obtain \((AB)\):
\[
\begin{pmatrix}
A \\ B
\end{pmatrix} = \frac{1}{\Delta} \left( \sum w_i x_i^2 \sum w_i y_i - \sum w_i x_i \sum w_i x_i y_i \right)
\]
where \( \Delta \) is defined as before. You can check by setting \( w_i = 1 \) that you recover the results of the unweighted least squares method previously covered.

For the weighted least squares, we conclude that the coefficients \( A \) and \( B \) from the fitting procedures should be calculated according to the formulae:
\[
A = \frac{\sum w_i x_i^2 \sum w_i y_i - \sum w_i x_i \sum w_i x_i y_i}{\Delta},
\]
\[
B = \frac{\sum w_i \sum w_i x_i y_i - \sum w_i x_i \sum w_i y_i}{\Delta}.
\]

5.7.2. Fitting Exponentials using a Straight Line Model. The exponential function
\[
y(x) = Ae^{Bx}
\]
does not appear to be amenable to the technique of linear least squares because it is not linear in the parameters \( A \) and \( B \). Such exponential functions are ubiquitous in nature. They arise as solutions of differential equations of the type:
\[
dy/dt = \lambda y \quad \Rightarrow \quad y(t) = y(0)e^{\lambda(t-t_0)},
\]
as you can easily check by substitution of \( y(t) \) into the differential equation.
As an example, the intensity \( I \) of a radiation beam penetrating a distance \( x \) through a shield obeys the equation (see Fig. 5.8)
\[
I(x) = I_0e^{-\mu x}
\]
\( I_0 \): intensity of the incident beam
\( \mu \): absorption coefficient (a property of the shield material)
Another example is an \( RC \) circuit (Fig. 5.9). The charge \( Q \) accumulated on capacitor \( C \) drains away exponentially fast when the capacitor is connected to a resistor. The time-dependence of the charge is described by:
\[
Q(t) = Q_0e^{-\lambda t}
\]
\( Q_0 \): initial charge (at time \( t = 0 \))
\( \lambda = 1/RC \): inverse time constant
5.7. Weighted Least Squares

![Graph of beam intensity decay](image)

Figure 5.8. Decay of beam intensity as function of depth of penetration into a material. Attenuation is usually exponential with distance.

![RC circuit](image)

Figure 5.9. RC circuit.

\( R \): resistance  
\( C \): capacitance

Fortunately, these models can be treated by linear least squares if we take the log of the equation:

\[
\log y = \log A + Bx
\]

and fit a straight line to the equation \((z_i = \log y_i)\):

\[
z_i = \log A + Bx_i
\]

If the errors \(\sigma_y\) in \(y\) are identical at each point, then for a function of the form \(z = \log y\), \(\sigma_z\) depends on \(y\) according to \(\sigma_z = \left|\frac{dz}{dy}\right|\sigma_y = \frac{\sigma_y}{y}\). We can then use the weighted least squares method.

5.7.3. Fitting a Polynomial. We can derive least squares formulae for higher order polynomial models (see Fig. 5.10)

\[
y(x; A, B, C, \ldots, H) = A + Bx + Cx^2 + \cdots + Hx^{p-1}
\]

For example, the height of a falling body should obey the equation

\[
y(t; y_0, v_0, g) = y_0 + v_0t - \frac{1}{2}gt^2
\]
5. Data Fitting

Figure 5.10. In general, a polynomial can be fit to most smooth curves. The challenge is to choose a fitting model that has the lowest number of parameters possible while fitting the data well.

where

$y_0, v_0$: initial height and velocity, respectively

$g$: acceleration due to gravity.

Consider the model:

$$y(x; A, B, C) = A + Bx + Cx^2,$$

which is still considered a “linear model” because linear refers to linearity in the fitting parameters $A$, $B$ and $C$, which this function fulfills. The corresponding chi-square function is:

$$\chi^2 = \sum_{i=1}^{n} \frac{(y_i - A - Bx_i - Cx_i^2)^2}{\sigma_y^2}.$$

To obtain the coefficients $A$, $B$ and $C$ we invoke the principle of maximum likelihood. The likelihood function

$$L(A, B, C | y_1, \ldots, y_n) = p(y_1, \ldots, y_n) \propto e^{-\chi^2/2},$$

is viewed as a function of the fitting parameters $A, B, C$. Minimization is performed with respect to these parameters. Setting $\partial \chi^2 / \partial A = 0$, $\partial \chi^2 / \partial B = 0$ and $\partial \chi^2 / \partial C = 0$:

$$An + B \sum_{i=1}^{n} x_i + C \sum_{i=1}^{n} x_i^2 = \sum_{i=1}^{n} y_i$$

$$A \sum_{i=1}^{n} x_i + B \sum_{i=1}^{n} x_i^2 + C \sum_{i=1}^{n} x_i^3 = \sum_{i=1}^{n} x_i y_i$$

$$A \sum_{i=1}^{n} x_i^2 + B \sum_{i=1}^{n} x_i^3 + C \sum_{i=1}^{n} x_i^4 = \sum_{i=1}^{n} x_i^2 y_i$$

These 3 equations in 3 unknowns can be solved to yield values of $A$, $B$ and $C$ in terms of the data.

5.7.4. General Linear Function of the Parameters. The method works for functions $y = f(x)$ which depend linearly on the coefficients $A$, $B$, \ldots
For example,
\[ y(x; A, B) = A \sin(x) + B \cos(x) \]
depends linearly on \( A \) and \( B \). Likewise, any function of the form:
\[ y(x; A, B, \ldots, H) = Af(x) + Bg(x) + \cdots + Hk(x) \]
where \( f, g, \ldots, k \) are known functions.

5.7.5. Multiple Regression. Many problems require 2 or more variables. An example is the ideal gas law \( PV = Nk_B T \) for fixed \( N \), which expresses the relationship between pressure, volume and temperature, i.e. \( P = f(V, T) \).

Suppose that
\[ z(x, y; A, B, C) = A + Bx + Cy \]
and we measure the data points (set of triples)
\[ \{(x_i, y_i, z_i)\}, \quad i = 1, \ldots, n \]
\( z_i \): all have the same uncertainty (\( \sigma_z \))
\( x_i, y_i \): are assumed to be "exact" (negligible uncertainty)

Maximum likelihood yields:
\[
\begin{align*}
An + B \sum_{i=1}^{n} x_i + C \sum_{i=1}^{n} y_i &= \sum_{i=1}^{n} z_i \\
A \sum_{i=1}^{n} x_i + B \sum_{i=1}^{n} x_i^2 + C \sum_{i=1}^{n} x_i y_i &= \sum_{i=1}^{n} x_i z_i \\
A \sum_{i=1}^{n} y_i + B \sum_{i=1}^{n} x_i y_i + C \sum_{i=1}^{n} y_i^2 &= \sum_{i=1}^{n} y_i z_i
\end{align*}
\]
and we solve for \( A, B \) and \( C \) to get the best fit (in the least squares sense).

5.7.6. Linear vs Nonlinear Dependence on Parameters. So far we have discussed the case of “linear least squares”. By linear, we mean that the model is a linear function of the model’s parameters. For example:
\[
\begin{align*}
y(x; A, B) &= A + Bx \\
y(x; A, B, C) &= A + Bx + Cx^2 \\
y(x; A, B, C, D, E) &= A + Bx + Cz + Dz^2 + Exz
\end{align*}
\]
are all linear models. In general, a linear model is of the form:
\[ y(x; \beta) = \sum_{i=1}^{n} \beta_i f_i(x) \]
where each term contains the first power of the $\beta_i$’s. The functions $f_i(x)$ can be nonlinear in $x$. In matrix form, $y(x; \vec{\beta}) = \vec{\beta}^T \vec{f}(x)$, where $\vec{\beta}^T$ is the row vector $\vec{\beta}^T = (\beta_1, \ldots, \beta_p)$ and $\vec{f}$ is the column vector $(f_1, \ldots, f_p)^T$ (the subscript $T$ denotes “transpose”). For example,

$$y(x; A, B) = A \cos(x) + B \log(x)$$

is still considered a “linear least squares” problem because $A$ and $B$ show up in first power. On the other hand, $y(x; A, B) = A \cos(Bx)$ is not linear in $B$ because $B$ shows up in even powers up to infinity. Similarly, $y(x; A) = A + A^2 x$ is nonlinear because of $A^2$. If the dependence on the parameters is linear, we can use the maximum likelihood technique to obtain estimates of the parameters in terms of the data. This is called “linear least squares”.

Non-Linear Least Squares Optimization

If the model is nonlinear in the fitting parameters we must use iterative techniques. In many cases, the best we can do is “guess” the initial parameter values and evolve them over time until we reach a satisfactory solution. Our focus is on the function $\chi^2(\vec{\beta})$ because $\chi^2$ measures the difference between data and a model, data fitting amounts to minimizing $\chi^2$:

$$\min_{w.r.t. (\beta_1, ..., \beta_p)} \chi^2(\beta_1, ..., \beta_p) = \chi^2(\vec{\beta}_1, ..., \vec{\beta}_p)$$

with respect to the parameters $\beta_1, ..., \beta_p$. These are the fitting parameters we previously denoted $A, B, ...$. These parameters span a multidimensional space. The parameters of $\chi^2$ are sometimes subjected to constraints.

If the data is fitted to a function $f(x; \beta_1, ..., \beta_p)$ that is linear in the parameters, we can use linear least squares. In other cases ($f$ is nonlinear in the $\beta_i$'s) we must resort to iterative techniques. The iterative approach consists of:

- Start with initial guess $\{\beta_i^0\}$.
- Iterate until $\chi^2$ is minimized, i.e. obtain a new set of coefficients $\{\beta_i^s\}$ from the previous ones $\{\beta_i^{s-1}\}$ using some suitable rule. (Very often, the iterations stop when $\chi^2$ no longer changes appreciably.)
- Final $\{\beta_i^{\text{final}}\}$'s yield a global minimum of $\chi^2$. At least, we hope that the minimum we reached is a global minimum.
6.1. Newton-Raphson method

Newton-Raphson is a method for finding successively better approximations to the roots (or zeros) of a real-valued function \( f(x) \). These zeros are solutions to the equation \( f(x) = 0 \). When applied to the function \( f'(x) \), it can be used to find extrema of \( f \).

6.1.1. Finding Zeros of a Function. We can find the zeros of a function, \( \{x|f(x) = 0\} \), even when \( x \) is defined implicitly. For example, suppose you are asked to find the maximum likelihood estimator of \( \alpha \), a parameter of the gamma distribution. The likelihood function for a set of gamma distributed random variables is:

\[
L(\alpha, \beta) \equiv p(x_1, \ldots, x_n; \alpha, \beta) = \frac{\beta^n \alpha}{(\Gamma(\alpha))^n} \prod_{i=1}^n x_i^{\alpha-1} e^{-\beta \sum_{i=1}^n x_i}.
\]

where \( x_i \) are the experimentally measured data and \( \Gamma(\cdot) \) is the gamma function. Taking the log of \( L \) and differentiating with respect to \( \beta \) gives the maximum likelihood estimator:

\[
\hat{\beta} = \frac{\hat{\alpha}}{n \sum_{i=1}^n x_i}.
\]

However, the derivative with respect to \( \alpha \) gives

\[
\log(\hat{\alpha}) - \frac{\Gamma'(\hat{\alpha})}{\Gamma(\hat{\alpha})} = \log \left( \frac{1}{n} \sum_{i=1}^n x_i \right) - \frac{1}{n} \sum_{i=1}^n \log x_i,
\]

an expression that is highly nonlinear in \( \hat{\alpha} \). We cannot solve for \( \alpha \) by writing this equation in the form \( \hat{\alpha} = h(\{x_i\}) \), for some function \( h \) of the data points \( \{x_i\} \). Instead, we can write this equation in the form \( f(\hat{\alpha}; \{x_i\}) = 0 \) and use the Newton-Raphson method to solve for \( \hat{\alpha} \) in terms of the \( \{x_i\} \).

The Newton-Raphson algorithm consists of:

- Choose a starting point, \( x_1 \).
- Approximate \( f(x_1) \) near \( x_1 \).

\[
f(x_1 + h) \equiv 0 \approx f(x_1) + f'(x_1)h.
\]

so that

\[
f(x_1) + f'(x_1)h \approx 0
\]

and therefore

\[
h \approx -\frac{f(x_1)}{f'(x_1)}.
\]
6.2. Finding Extrema of a Function

If we want to find extrema of a function \( f(x) \), we can apply the Newton-Raphson method to the function \( f'(x) \) instead of \( f(x) \), as shown in Fig. 6.2. The update rule is:

\[
x_{s+1} = x_s - \frac{f'(x_s)}{f''(x_s)}.
\]

In other words, Taylor expanding the gradient of \( f \) around the point \( x_s \) and setting equal to zero gives the update rule for extrema of a function:

\[
(6.1) \quad f'(x_s + h) = f'(x_s) + h \cdot f''(x_s) = 0 \quad \Rightarrow \quad h = -\frac{f'(x_s)}{f''(x_s)},
\]

which leads to

\[
x_{s+1} = x_s - \frac{f'(x_s)}{f''(x_s)}.
\]
6.3. Gradient (Steepest) Descent Method

∇χ^2 lies along the direction where χ changes most rapidly (Fig. 6.3). Consequently, we may construct an update rule as

\[ \tilde{β}_{s+1} = \tilde{β}_s - λ∇χ^2(\tilde{β}_s). \]

where λ > 0 is a scaling parameter which should not be too large (to avoid overshoot) and not too small (so it converges reasonably fast) while \( \tilde{β}_s \) is a column vector containing all \( p \) parameters at the \( s \)-th iteration. (The negative sign is used because \( ∇χ^2 \) points in the direction of steepest increase of \( χ^2 \) and we want the direction of steepest decrease.) \( ∇χ^2 \) can be computed analytically if the functional form of \( χ^2 \) (hence, the model) is known.

If we don’t have an analytical formula available for the gradient, we can approximate \( (∇χ^2)_j = \frac{∂χ^2}{∂β_j} \) numerically by finite differences:

\[ \frac{χ^2(β_1, ..., β_j-1, β_j + δβ_j, β_{j+1}, ..., β_p) - χ^2(β_1, ..., β_{j-1}, β_j, β_{j+1}, ..., β_p)}{δβ_j}, \]

where \( δβ_j \) is a small step along the \( j \)-th direction.

This method moves toward the minimum because the directional derivative of \( f \) along \( \hat{y} \) evaluated at the point \( \vec{a} \)

\[ \frac{df}{dy}(\vec{a}) \equiv ∇f(\vec{a}) \cdot \hat{y} = ∥∇f(\vec{a})∥ \cos \theta \]

is largest when \( \hat{y} \) points along (parallel to) the vector \( ∇f(\vec{a}) \) (i.e. when \( \theta = 0 \)). The function \( f \) undergoes its maximum rate of change in the direction of \( ∇f(\vec{a}) \) (Fig. 6.4). Thus, the gradient \( ∇f(\vec{a}) \) is a vector that points along the direction of steepest increase of \( f \) (at the point \( \vec{a} \)).
6.4. Newton Method

Near a minimum \( \tilde{\beta_s} \), \( \chi^2 \) looks parabolic whereas at the minimum we have \( \nabla \chi^2 = 0 \). Let us proceed as we did for the Newton-Raphson method (see equation 6.1) by Taylor expanding the gradient of \( \chi^2 \) near the minimum:

\[
\nabla \chi^2(\tilde{\beta_s} + \vec{h}) = \nabla \chi^2(\tilde{\beta_s}) + \vec{h} \cdot \nabla \nabla \chi^2(\tilde{\beta_s}) = 0,
\]

where \( \nabla \nabla \chi^2(\tilde{\beta_s}) \) is the \( p \times p \) Hessian matrix:

\[
H_s \equiv \nabla \nabla \chi^2(\tilde{\beta_s}) = \begin{bmatrix}
\frac{\partial^2 \chi^2}{\partial \beta_1^2} & \cdots & \frac{\partial^2 \chi^2}{\partial \beta_1 \partial \beta_p} \\
\vdots & \ddots & \vdots \\
\frac{\partial^2 \chi^2}{\partial \beta_p \partial \beta_1} & \cdots & \frac{\partial^2 \chi^2}{\partial \beta_p^2}
\end{bmatrix}.
\]

Solving for \( \vec{h} \) gives

\[
\vec{h} = -H_s^{-1}\nabla \chi^2(\tilde{\beta_s}).
\]
The update rule \( \beta_{s+1} = \beta_s + h \) is:

\[
\beta_{s+1} = \beta_s - H_s^{-1} \nabla \chi^2(\beta_s).
\]

This requires calculating and inverting a Hessian matrix. If \( H_s = 1/\lambda \) (\( 1 \) is a \( p \times p \) unit matrix) this method is equivalent to the steepest descent. \( H_s \) is also called the curvature matrix because second partial derivatives, \( \partial^2 \chi^2 \), are indicative of the curvature of the function (\( \chi^2 \) here) along each direction \( i \).

### 6.5. Levenberg Method

The steepest descent method works best on flat terrain (away from the minimum) whereas the Newton method works best near the minimum, which is well approximated by a parabola. Levenberg proposed an improved update rule which interpolates between the two methods:

\[
\beta_{s+1} = \beta_s - (H_s + \lambda 1)^{-1} \nabla \chi^2(\beta_s),
\]

which is a modified Newton method. \( 1 \) is a \( p \times p \) unit matrix. The term \( \lambda 1 \) is a regularization term. When \( \lambda \) is large (denoting \( \lambda^{-1} = \tilde{\lambda} \)), \( \|H_s\| \ll \|\lambda 1\| \),

\[
\beta_{s+1} = \beta_s - (H_s + \lambda 1)^{-1} \nabla \chi^2(\beta_s) \approx \beta_s - (\lambda 1)^{-1} \nabla \chi^2(\beta_s) = \beta_s - \tilde{\lambda} \nabla \chi^2(\beta_s),
\]

we recover the steepest descent method. On the other hand, when \( \lambda \) is small,

\[
\|H_s\| \gg \|\lambda 1\|,
\]

and

\[
\beta_{s+1} = \beta_s - (H_s + \lambda 1)^{-1} \nabla \chi^2(\beta_s) \approx \beta_s - (H_s)^{-1} \nabla \chi^2(\beta_s) = \beta_s - H_s^{-1} \nabla \chi^2(\beta_s),
\]

we recover the Newton method. Far from the minimum, we want to use large values of \( \lambda \) whereas close to the minimum we want to use small values of \( \lambda \). The parameter \( \lambda \) is adjusted at each iteration. We stop iterating when \( \chi^2 \) does not change appreciably.

### 6.6. Gauss-Newton Method

In the Newton method the update rule is:

\[
\beta_{s+1} = \beta_s - H_s^{-1} \nabla \chi^2(\beta_s) \quad \text{(Newton)}
\]

It is a second-order method because it involves computing the Hessian (matrix of second partial derivatives). We must also invert the Hessian matrix. The Gauss-Newton method is also a “second order” method, but a simpler form of the Hessian matrix is used. Let

\[
\chi^2 = \sum_{i=1}^n \tilde{R}_i^2, \quad \tilde{R}_i = \frac{y_i - y(x_i; \tilde{\beta})}{\alpha_i} \sim N(0, 1)
\]

\(^1\)Here \( \|A\| \) denotes the norm of the matrix \( A \). Any suitable norm can be used. For example, it can be the largest of all matrix entries: \( \|A\| = \max_{ij} |A_{ij}| \).
where \( n \) is the total number of data points. Define the \( n \times p \) Jacobian matrix:

\[
J_s \equiv J(\vec{\beta}_s) = \begin{bmatrix}
\frac{\partial \tilde{R}_1}{\partial \beta_1} & \cdots & \frac{\partial \tilde{R}_1}{\partial \beta_p} \\
\vdots & \ddots & \vdots \\
\frac{\partial \tilde{R}_n}{\partial \beta_1} & \cdots & \frac{\partial \tilde{R}_n}{\partial \beta_p}
\end{bmatrix}.
\]

Now, the gradient of \( \chi^2 \) is

\[
(\nabla \chi^2)_j \equiv \frac{\partial \chi^2}{\partial \beta_j} = \frac{\partial}{\partial \beta_j} \sum_{i=1}^{n} \tilde{R}_i^2 = 2 \sum_i \frac{\partial \tilde{R}_i}{\partial \beta_j} \tilde{R}_i,
\]

which can be written compactly as \( \nabla \chi^2 = \nabla \sum_{i=1}^{n} \tilde{R}_i^2 = 2J^T \tilde{R} \), where \( \tilde{R} \) is the vector of normalized residuals (one entry for each of the \( n \) data points):

\[
\tilde{R} \equiv \begin{bmatrix}
\tilde{R}_1 \\
\tilde{R}_2 \\
\vdots \\
\tilde{R}_n
\end{bmatrix}.
\]

The Hessian matrix is therefore:

\[
(H_s)_{jk} \equiv \frac{\partial^2 \chi^2}{\partial \beta_j \partial \beta_k} = \frac{\partial^2}{\partial \beta_j \partial \beta_k} \sum_{i=1}^{n} \tilde{R}_i^2 = 2 \sum_i \frac{\partial \tilde{R}_i}{\partial \beta_j} \frac{\partial \tilde{R}_i}{\partial \beta_k} + 2 \sum_i \tilde{R}_i \frac{\partial^2 \tilde{R}_i}{\partial \beta_j \partial \beta_k},
\]

where the last term is neglected since \( \tilde{R}_i \) is small (\( O(1) \)) and normally distributed with mean 0 and variance 1. This yields \( H_s \approx 2J_s^T J_s \), which is easier to evaluate than the full Hessian. This has the advantage of requiring fewer steps to compute. The final update rule is:

\[
\vec{\beta}_{s+1} = \vec{\beta}_s - (2J_s^T J_s)^{-1} \nabla \chi^2(\vec{\beta}_s) \quad \text{(Gauss-Newton)}
\]

The disadvantage of second-order methods is they don’t work well outside parabolic surfaces (hence the development of the Levenberg method and Marquardt-Levenberg; the latter is described in the next section).

### 6.7. Marquardt-Levenberg Method

We recall the Levenberg update rule:

\[
\vec{\beta}_{s+1} = \vec{\beta}_s - (H_s + \lambda I)^{-1} \nabla \chi^2(\vec{\beta}_s), \quad \text{(Levenberg)}
\]

which interpolates between steepest descent and the Newton method by way of the regularization term, \( \lambda I \). Indeed, for large \( \lambda \) we have the steepest descent method and for small \( \lambda \) we recover the Newton method. Far from the minimum, we want to use large values of \( \lambda \) whereas close to the minimum
we want to use small values of $\lambda$.

An improvement over the Levenberg method was proposed by Marquardt in the form of a modified update rule:

$$
\tilde{\beta}_{s+1} = \tilde{\beta}_s - (H_s + \lambda \text{diag}[H_s])^{-1} \nabla \chi^2(\tilde{\beta}_s).
$$

(Marquardt-Levenberg)

where $\text{diag}[H_s]$ is the matrix $H_s$ where all entries have been zeroed out except those along the diagonal:

$$
\text{diag}[H_s] = \text{diag} \begin{bmatrix}
H_{11} & H_{12} & \ldots & H_{1p} \\
H_{21} & H_{22} & \ldots & H_{2p} \\
\vdots & \vdots & \ddots & \vdots \\
H_{p1} & \ldots & \ldots & H_{pp}
\end{bmatrix} = \begin{bmatrix}
H_{11} & 0 & \ldots & 0 \\
0 & H_{22} & \ldots & 0 \\
\vdots & \vdots & \ddots & \vdots \\
0 & \ldots & \ldots & H_{pp}
\end{bmatrix}.
$$

The Marquardt-Levenberg method is equivalent to modifying the Hessian to $H_{jj} \rightarrow H_{jj}(1 + \lambda)$ and $H_{jk} \rightarrow H_{jk}$ ($j \neq k$). For large $\lambda$, the matrix $H_s + \lambda \text{diag}[H_s]$ is said to be “diagonally dominant”, i.e. it has the form

$$
H_s + \lambda \text{diag}[H_s] \approx (1 + \lambda) \begin{bmatrix}
H_{11} & 0 & \ldots & 0 \\
0 & H_{22} & \ldots & 0 \\
\vdots & \vdots & \ddots & \vdots \\
0 & \ldots & \ldots & H_{pp}
\end{bmatrix}.
$$

(large $\lambda$)

where $H_{ii} = \frac{\partial^2 \chi^2}{\partial \beta_i \partial \beta_i}$ is the curvature of $\chi^2$ surface along $i$-th direction. The inverse of a diagonal matrix involves taking the inverse of every diagonal element:

$$
(H_s + \lambda \text{diag}[H_s])^{-1} \approx \frac{1}{1 + \lambda} \begin{bmatrix}
\frac{1}{H_{11}} & 0 & \ldots & 0 \\
0 & \frac{1}{H_{22}} & \ldots & 0 \\
\vdots & \vdots & \ddots & \vdots \\
0 & 0 & \ldots & \frac{1}{H_{pp}}
\end{bmatrix}.
$$

This matrix then multiplies the column vector $\nabla \chi^2(\tilde{\beta}_s)$ to yield the update rule:

$$
\tilde{\beta}_{s+1} = \tilde{\beta}_s - (H_s + \lambda \text{diag}[H_s])^{-1} \nabla \chi^2(\tilde{\beta}_s),
$$
where for large $\lambda$, this approximates to:

$$
\approx \vec{\beta}_s - (1 + \lambda)^{-1} \begin{bmatrix}
(\vec{g}_s)_1 \\
(\vec{g}_s)_2 \\
\vdots \\
(\vec{g}_s)_p \\
H_{s1} \\
\vdots \\
H_{sp}
\end{bmatrix} = \vec{\beta}_s - (1 + \lambda)^{-1} \begin{bmatrix}
\frac{\partial \chi^2}{\partial \beta_1^2} \\
\frac{\partial \chi^2}{\partial \beta_2^2} \\
\vdots \\
\frac{\partial \chi^2}{\partial \beta_p^2}
\end{bmatrix}.
$$

We used the shorthand notation:

$$
\partial_i \chi^2 \equiv \frac{\partial \chi^2}{\partial \beta_i},
$$

and

$$
\partial_{ii}^2 \chi^2 \equiv \frac{\partial^2 \chi^2}{\partial \beta_{ii}^2}.
$$

This update rule looks like the gradient (steepest) descent method, except that each entry has been scaled in such a way that large steps are made in the direction of low curvature (flat terrain) and small steps in the direction with high curvature (steep incline). Levenberg-Marquardt is considered one of the best “local optimization” algorithms and is widely used in applications. In MATLAB, it is implemented in the function \texttt{lsqnonlin}.

6.7.1. Adjusting $\lambda$ in Levenberg and Marquardt-Levenberg methods. The parameter $\lambda$ is adjusted at each iteration. We stop iterating when $\chi^2$ does not change appreciably. $H_s$ is called the curvature matrix. Here is a possible implementation of the Levenberg-Marquardt method (or Levenberg method):

- Pick initial guess for set of fitted parameters $\vec{\beta}$.
- Compute $\chi^2(\vec{\beta})$.
- Pick a value for $\lambda$, say $\lambda = 0.1$.
- (*) Solve for $\delta \vec{\beta} = -(H_s + \lambda \text{diag}[H_s])^{-1} \nabla \chi^2(\vec{\beta}_s)$ (Marquardt-Levenberg update rule; for the Levenberg method, use the Levenberg update rule instead) and evaluate $\chi^2(\vec{\beta} + \delta \vec{\beta})$.
- If $\chi^2(\vec{\beta} + \delta \vec{\beta}) \geq \chi^2(\vec{\beta})$, we increase $\lambda$ by a factor of 10 and go back to (*).
- If $\chi^2(\vec{\beta} + \delta \vec{\beta}) < \chi^2(\vec{\beta})$, we decrease $\lambda$ by a factor of 10, update the trial solution $\vec{\beta} \to \vec{\beta} + \delta \vec{\beta}$ and go back to (*). [In other words, if the total distance error of the updated parameters is less than the previous one, the updated parameters are taken to be the current parameters and $\lambda$ is decreased.]
• Stopping criterion: changes in parameters that yield changes in $\chi^2$ by an amount $\ll 1$ are not statistically meaningful.

• When finished, use the Hessian to compute the errors in the fitted parameters (see section 6.8).

The parameter $\lambda$ can be initialized to be large so that first updates are small steps in the steepest-descent direction. If an iteration happens to result in a worse approximation, $\lambda$ is increased. As the solution improves, $\lambda$ is decreased, the Levenberg-Marquardt method approaches the Gauss-Newton method, and the solution typically accelerates to the local minimum.

For more information about this algorithm see:


6.7.2. Confidence regions. If we plot contour lines of equal $\chi^2$ (see Fig. 6.5), the interior region of these contours can be associated with the likelihood of a set of fitted parameters (random vector) lies within that contour.

6.7.3. Local optimization techniques: summary of update rules. The algorithms covered so far for nonlinear optimization are called "local
optimization” techniques, because they are designed to search for the nearest minimum of $\chi^2$. This minimum may not necessarily be the global minimum of the $\chi^2$ surface. In subsequent lectures we will look at global optimization techniques. We summarize in Table 6.1 the update rules derived so far, for these local optimization techniques.

### Table 6.1. Various update rules for non-linear local optimization.

<table>
<thead>
<tr>
<th>Method</th>
<th>Update rule (displacement at time $s$)</th>
</tr>
</thead>
<tbody>
<tr>
<td>gradient descent</td>
<td>$\vec{h} = -\lambda \nabla \chi^2(\vec{\beta}_s)$</td>
</tr>
<tr>
<td>Newton method</td>
<td>$\vec{h} = -(\mathbf{H}_s)^{-1} \nabla \chi^2(\vec{\beta}_s)$</td>
</tr>
<tr>
<td>Gauss-Newton</td>
<td>$\vec{h} = -\frac{1}{2} (\mathbf{J}_s^T \mathbf{J}_s)^{-1} \nabla \chi^2(\vec{\beta}_s)$</td>
</tr>
<tr>
<td>Levenberg</td>
<td>$\vec{h} = -((\mathbf{H}_s + \lambda 1)^{-1} \nabla \chi^2(\vec{\beta}_s)$</td>
</tr>
<tr>
<td>Levenberg-Marquardt</td>
<td>$\vec{h} = -(\mathbf{H}_s + \lambda \text{diag}[\mathbf{H}_s])^{-1} \nabla \chi^2(\vec{\beta}_s)$</td>
</tr>
</tbody>
</table>

6.8. Errors in the fitting parameters: the covariance matrix

The errors in the fitted parameters $\vec{\beta}$ can be extracted from the main diagonal of the covariance matrix (omitting the subscript $s$ in $\vec{\beta}_s$ momentarily to avoid cluttering the notation):

$$
cov(\vec{\beta}, \vec{\beta}) = \begin{bmatrix}
\text{var}(\beta_1) & \text{cov}(\beta_1, \beta_2) & \ldots & \text{cov}(\beta_1, \beta_p) \\
\text{cov}(\beta_2, \beta_1) & \text{var}(\beta_2) & \ldots & \text{cov}(\beta_1, \beta_p) \\
\vdots & \vdots & \ddots & \vdots \\
\text{cov}(\beta_p, \beta_1) & \text{cov}(\beta_p, \beta_2) & \ldots & \text{var}(\beta_p)
\end{bmatrix}
$$

whereas the off-diagonal elements describe possible relationships (e.g. such as redundancy) among the fitting parameters. But how do we obtain the matrix $cov(\vec{\beta}, \vec{\beta})$? It can be shown (see Chapter 8 for proof) that the covariance matrix can be obtained from the Hessian:

$$
cov(\vec{\beta}_s, \vec{\beta}_s) = 2 \mathbf{H}_s^{-1}
$$

For this to work, of course, the matrix $\mathbf{H}_s$ needs to be invertible (non-singular). This may not always be the case. (The subscript $s$ here indicates that covariances can be monitored in real-time, at every time step of the iterative optimization process.)

6.9. Problems

**Problem 71.** Regarding the problem of data fitting, where we need gradients, Jacobians, etc. (review lecture notes for the definitions of $\nabla \chi^2(\vec{\beta}_s)$, $\mathbf{J}_s$, etc.). (a) Take $y(x) = Ax + B$ as the fitting model. Calculate $\nabla \chi^2(\vec{\beta}_s)$,
6. Non-Linear Least Squares Optimization

\( \mathbf{J}_s, \mathbf{H}_s \) and the covariance matrix.

(b) Let the fitting model be given by \( y(x) = A + Bx + C \cos(Dx) \). Calculate \( \nabla \chi^2(\vec{\beta}_s) \) and \( \mathbf{J}_s \).

(c) The model is \( y(x) = A \exp(-Bx) + C \). Calculate \( \nabla \chi^2(\vec{\beta}_s), \mathbf{J}_s \) and \( \mathbf{H}_s \).

**Solution.** Let’s do the Jacobian here for (i), the linear model. From the definition of the Jacobian, this is a \( n \times p \) matrix of first partial derivatives of the normalized residuals. The rows are the data points and the columns are the fitting parameters (here, \( A \) and \( B \)):

\[
\mathbf{J}_s = \begin{bmatrix}
\frac{\partial \tilde{R}_1}{\partial A} & \frac{\partial \tilde{R}_1}{\partial B} \\
\vdots & \vdots \\
\frac{\partial \tilde{R}_n}{\partial A} & \frac{\partial \tilde{R}_n}{\partial B}
\end{bmatrix}
\]

where \( \tilde{R}_i = \frac{y_i - y(x_i; A, B)}{\alpha_i} \) and \( y(x_i; A, B) = Ax_i + B \). Thus,

\[
\mathbf{J}_s = \begin{bmatrix}
\frac{x_1}{\alpha_1} & \frac{1}{\alpha_1} \\
\vdots & \vdots \\
\frac{x_n}{\alpha_n} & \frac{1}{\alpha_n}
\end{bmatrix}
\]

**Problem 72.** The following model is to be used for data fitting

\[
y(x) = 40B \log x^2 + 4A \cos x + 2A
\]

where \( A \) and \( B \) are parameters to be determined from the experimental data. In terms of this model, find the Jacobian, Hessian and covariance matrices and explain their role/purpose. Indicate which entries of the covariance matrix measure the redundancy between the fitting parameters.
Global Optimization

The algorithms discussed so far can only take us to the nearest minimum. Once a minimum has been reached, there is no built-in mechanism to escape the minimum. If this minimum is not a global minimum, the solution obtained is not optimal. A general $\chi^2(\vec{\beta})$ surface:

$$\chi^2(\beta_1, \ldots, \beta_p) = \sum_{i=1}^{n} \frac{(y_i - y(x_i; \vec{\beta}))^2}{\sigma_i^2}.$$  

(textbook writes $\alpha_i$ instead of $\sigma_i$)

could potentially have several minima, as shown in Fig. 7.1 for the 2D case $\vec{\beta} = (\beta_1, \beta_2)$.

![Figure 7.1.](image)

Figure 7.1. Global optimization aims to find the global minimum of the $\chi^2$ surface.
The reason that gradient (or Hessian) based methods failed to see the global minimum is because they only possess knowledge of the first and second derivative of $\chi^2$ near the point $\vec{\beta}$. This local neighborhood of $\vec{\beta}$ does not extend very far into the parameter space.

Global optimization algorithms tend to rely on the use of multiple random initial guesses, or on the use of evolution steps that may include some element of randomness, or a combination of both. These random elements are essential in order to “push” the search toward regions of the parameter space ($\vec{\beta}$) that gradient-based techniques would otherwise not be able to reach. While global optimization schemes may or may not also use gradient-based searches as part of the optimization strategy, the “global” nature of the algorithm generally owes to the addition of some element of randomness.

Here we will cover two such global methods: the simulated annealing (Metropolis) method and the genetic algorithm. You should, however, be aware that there are many additional methods in use: quantum annealing, stochastic tunneling, tabu search, reactive search optimization, stochastic gradient descent, graduated optimization, ant colony optimization, cross-entropy method, harmony search, particle swarm optimization, intelligent water drops and parallel tempering. Each method has its advantages and disadvantages. Another ideal setting for these global optimization methods is in cases where derivative information is not available.

### 7.1. The Metropolis Algorithm (Simulated Annealing)

The Metropolis algorithm (Kirkpatrick, 1983) - also known as Metropolis Monte-Carlo - is a version of simulated annealing (SA) that utilizes the so-called Metropolis criteria [due to Metropolis (1953) and Hastings (1970)]. It is inspired by the annealing process in metallurgy. Annealing involves heating and controlled cooling of a metal or alloy to produce a high quality crystalline lattice. Heat increases thermal motion of the atoms and eases diffusional motion. The subsequent cooling causes the atoms to migrate to sites of local minimum energy, which usually translates into a lower amount of defects and a global energy minimum.

The algorithm can be summarized as follows:

1. Choose an initial configuration $\vec{\beta}$ for the fitting parameters. The starting parameters can be selected randomly.
2. Choose a starting temperature $T > 0$. Temperature should be high enough, as the goal of this algorithm is to cool (by annealing) over a long period of time.
3. Calculate $\chi^2(\vec{\beta})$ for this configuration.
4. Effect a random change $\delta \vec{\beta}$ to this configuration: $\vec{\beta}' = \vec{\beta} + \delta \vec{\beta}$. 
(5) Calculate the new energy $\chi^2(\vec{\beta}')$

(6) Accept the move with probability

$$P(\text{accept}) = \min \left(1, \exp\left(-\frac{\chi^2(\vec{\beta}') - \chi^2(\vec{\beta})}{T}\right)\right).$$

This step is called the Metropolis criterion. If the move is accepted, the new configuration $\vec{\beta}'$ is taken to be the current configuration, $\vec{\beta}$. If the move is rejected, the old configuration $\vec{\beta}$ is kept unchanged.

(7) Repeat steps 3-5 until convergence while lowering the temperature $T$.

A few words are in order. First, the temperature $T$ is dimensionless and is not a real temperature but a parameter that simulates the effects of temperature. The higher the temperature, the closer to 1 is the probability of acceptance, meaning that almost every state is accessible. At low temperatures, the exp(-·) (Boltzmann) factor is smaller and acceptance of the moves is less likely unless the energy of the system (as measured by $\chi^2$) decreases as a result of the new move.

Finite temperatures correspond to “thermal energy” supplied by a reservoir to the thermodynamic system. If the temperature is not lowered, the system will remain with the same average energy per unit volume and moves will continue to be accepted at the same rate. In order to find the “ground state” of the system the goal is to reach a point where the majority of moves are rejected. This can only happen at or near $T = 0$.

If the Boltzmann factor

$$\exp\left\{-\frac{\chi^2(\vec{\beta}') - \chi^2(\vec{\beta})}{T}\right\} = e^{-\Delta E/T}$$

has $\Delta E = \chi^2(\vec{\beta}') - \chi^2(\vec{\beta}) > 0$, the new move results in a higher energy configuration. In this case, rather than rejecting it, the move is accepted with probability $e^{-\Delta E/T} < 1$. This allows for the possibility of random jumps that could lift the system out of a local minimum. On the other hand, $\Delta E < 0$ corresponds to a move that lowers the energy of the system. In this case, $e^{-\Delta E/T} > 1$ and the move is accepted with probability 1 (always).

### 7.2. Accepting a Move With Probability $P$

What does it mean when we are asked to accept a move with probability $P$? For example,

$$P = e^{-\Delta E/T}. $$
Given a probability $P \in [0,1]$, regardless of the probability distribution it originates from, we may decide whether or not a random experiment occurs with probability $P$ as follows:

- Generate a uniformly distributed random number $R$ in the interval $R \in [0,1]$
- If $P > R$, accept the move (the event has occurred).
- If $P \leq R$, reject the move (the event did not occur).

The rationale for this reasoning is explained in the section below.

### 7.3. Sampling From a Distribution

In the Metropolis scheme and many other algorithms, we are asked to sample from a distribution, such as the exponential distribution. However, your random number generator may not be able to generate samples from the distribution of your choice. It is likely able to generate uniformly distributed samples $U([0,1])$ in the interval $[0,1]$. For example, the command `rand` in MATLAB will generate uniformly distributed random numbers:

```
>> help rand
rand Uniformly distributed pseudorandom numbers.
    R = rand(N) returns an N-by-N matrix containing pseudorandom values drawn
    from the standard uniform distribution on the open interval(0,1). rand(M,N)
    or rand([M,N]) returns an M-by-N matrix. rand(M,N,P,...) or
    rand([M,N,P,...]) returns an M-by-N-by-P-by-... array. rand returns a
    scalar. rand(SIZE(A)) returns an array the same size as A.
```

Fortunately, MATLAB has another command, `randn` which can generate random numbers sampled from a Gaussian distribution:

```
>> help randn
randn Normally distributed pseudorandom numbers.
    R = randn(N) returns an N-by-N matrix containing pseudorandom values drawn
    from the standard normal distribution. randn(M,N) or randn([M,N]) returns
    an M-by-N matrix. randn(M,N,P,...) or randn([M,N,P,...]) returns an
    M-by-N-by-P-by-... array. randn returns a scalar. randn(SIZE(A)) returns
    an array the same size as A.
```

But to my knowledge, there are no random number generators for other distribution functions. The problem also exists with many programming
languages such as C or FORTRAN, where random number generators exist but only for uniform distributions.

Luckily, a random number sampled from the uniform distribution, \( U([0,1]) \), can be used to generate random samples from any distribution function. Suppose that we want to generate Gaussian random numbers \( X \sim N(0,1) \) for a given random variable \( X \). The probability density is \( p_X(x) = \frac{1}{\sqrt{2\pi}} e^{-x^2/2} \).

In the figure below, we plot the cdf as the graph \((x,y(x))\):

\[
y(x) = P(X < x) = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{x} e^{-x'^2/2} \, dx' = \int_{-\infty}^{x} p_X(x') \, dx'
\]

in the case of a Gaussian density. The left-most plot shows a probability density function \( p_X(x) \) corresponding to some random variable \( X \). On the right, we have its cumulative distribution function \( y(x) = \int_{-\infty}^{x} p_X(x') \, dx' \).

We generate uniformly distributed samples \( y \) in the interval \([0,1]\) and invert \( y = y(x) \) to produce a sample \( x \) according to the desired distribution \( p_X(x) \). This is illustrated in Figure 7.2.

If we consider the \( y \) axis as a random variable \( Y \sim U([0,1]) \) which is sampled uniformly on the interval \([0,1]\), let us check that a uniform distribution for \( Y \) gives rise to the desired distribution for \( X \).

\[
P(Y < y(x)) = P \left( Y < \int_{-\infty}^{x} p_X(x') \, dx' \right) = \int_{0}^{\int_{-\infty}^{x} p_X(x') \, dx'} 1 \cdot dy = \int_{-\infty}^{x} p_X(x) \, dx = P(X < x).
\]

Therefore, a uniform distribution for \( Y \) implies picking \( X \) according to the density \( p_X(x) \).
7.4. Genetic Algorithms

Genetic algorithms (Holland, 1975) enable us to find global minima. In the early days, GAs have been applied to least squares curve fitting (Karr, 1991; Rogers 1991). Nowadays, GAs are applied to many different optimization problems in the physical and social sciences. Software packages such as MATLAB, IDL, Mathematica or Octave feature implementations of genetic algorithms (GA) that are relatively easy to use. One advantage of GA codes is they are inherently parallel and easily implemented on parallel hardware. Another advantage is they can handle large data sets. Their main disadvantage is they are slow. The general idea consists of generating candidate solutions and sending them to an evaluator for testing. If a candidate solution is not optimal, then the procedure is repeated. In genetic algorithms, the procedure is repeated iteratively over a large set of candidate solutions. Because this set can be large, a significant number of possible solutions can be tested simultaneously.

They are a type of parallel heuristic search method inspired by the laws of nature (genetics) that govern evolution of biological organisms. Each candidate solution is called an organism. A chromosome is a list of elements called genes. In the simplest case, an organism consists of a single chromosome (haploid), although there are cases when the organism consists of dual-strand chromosomes (diploid). Chromosomes usually consist of linear lists of genes. A gene can assume any of a number of values called alleles, which are taken from the base set. Generally, problem solutions are encoded as strings of alleles (most commonly, strings of 0’s and 1’s).

7.4.1. General idea. Let us look at an example strategy for a possible genetic algorithm. The algorithm below is purposely left vague so you can see the main steps involved and the parallel with evolution. In real implementations, the specifics of the algorithm strongly depend on the application. We will look at a specific example later.

1. Generate a large population of random chromosomes \( \{ \mathbf{\beta}_i \}, i = 1, \ldots, M \).
2. Each chromosome is assigned a fitness score \( F \) proportional to some goodness-of-fit parameter, e.g. \( \chi^2(\mathbf{\beta}) \). This fitness score \( F \) should increase as \( \chi^2 \) decreases.
3. Select the top fitness scores. These “parent” chromosomes, \( \{ \mathbf{\beta}_{\text{best}} \} \), will be used to breed the next generation (“offspring”).
4. Generate “offspring” from the parent chromosomes: \( \{ \mathbf{\beta}_i \}, i = 1, \ldots, M \).
7.4. Genetic Algorithms

(5) Introduce random changes in the genetic code in the form of crossover and mutations. Mutations are “tweaks” to the chromosome contents (e.g. flip a 0 to a 1). Crossover involves picking a gene at random and generating a new chromosome that consists of chromosome 1 up until that gene followed by the rest of chromosome 2, as indicated in Fig. 7.3.

(6) Repeat steps 2-5 until solution is found.

The inherently parallel nature of the algorithm is embodied in steps 2 and 4, where the GAs search large numbers of candidate solutions simultaneously. A decision is made only after all the candidate solutions have been generated. The objective function used by genetic algorithms is based on actual, problem-specific information, rather than auxiliary information, such as a gradient or a Hessian.

Step 3 is analogous to natural selection. Like its natural selection counterpart in biology, the selection operation selects pairs of highly fit organisms for mating. This focus toward the highly fit individuals is what drives genetic algorithms.

The genetic algorithm analogy to mating is called crossover. The crossover operation provides a mixing of the genes from the parents, and globally it mixes the genetic material of the whole population. It is the mixing of the genes, the stirring of the pot of genetic material, that gives robustness to the genetic algorithm.

The two organisms chosen by selection are combined to form a new individual with similarities to both parents. If the mixing is done carefully, then a large amount of genetic material can be tested. Although selection focuses on the genetic algorithm, it is crossover that adds variety.

The method based on the random selection of a single crossover point described above is the simplest implementation of crossover. More complicated implementations are possible. There is no specific guidelines as to what rule should be used. One could, for example, take the average of two chromosomes, their product, their difference or their sum. Which implementation is best suited depends on application.
7.5. Review of Calculus: Partial derivatives and Taylor’s theorem in several variables

Because nonlinear optimization methods make extensive use of partial derivatives, here we review partial derivatives and the Taylor’s theorem in multiple variables. We will show how to compute the partial derivatives of \( \frac{1}{r} \), where \( r = |r| \) and \( r \) has components \( r = (x, y, z) \). \( r \) is its Euclidean length:

\[
|r| \equiv r \equiv \sqrt{x^2 + y^2 + z^2}.
\]

7.5.1. Einstein summation convention. To keep the notation neat (uncluttered), we will use the Einstein summation convention. Whenever two indices are repeated in the same term, a summation is implied. For example, in the dot product of \( u = (u_x, u_y, u_z) \) and \( v = (v_x, v_y, v_z) \) we have:

\[
\mathbf{u} \cdot \mathbf{v} = u_\alpha v_\alpha \equiv \sum_{i=1}^{3} u_i v_i = u_x v_x + u_y v_y + u_z v_z.
\]

It is simpler to write \( u_\alpha v_\alpha \) than the entire summation.

7.5.2. Multivariate Taylor expansion. In 1D the Taylor expansion of \( f(x + h) \) at \( x \) is:

\[
f(x + h) = f(x) + f'(x)h + \frac{1}{2} f''(x)h^2 + \frac{1}{3!} f^{(3)}(x)h^3 + \ldots
\]

In \( n \)-D, a scalar-valued function \( f(x) \), where \( x \in \mathbb{R}^n \), is expanded as:

\[
f(x + h) = f(x) + \sum_{i=1}^{n} \frac{\partial f(x)}{\partial x_i} h_i + \frac{1}{2} \sum_{i=1}^{n} \sum_{j=1}^{n} \frac{\partial^2 f(x)}{\partial x_i \partial x_j} h_i h_j + \frac{1}{3!} \sum_{i=1}^{n} \sum_{j=1}^{n} \sum_{k=1}^{n} \frac{\partial^3 f(x)}{\partial x_i \partial x_j \partial x_k} h_i h_j h_k + \ldots
\]

7.5.3. Abbreviation for Partial Derivatives. We will often use the shorthand notation to abbreviate the notation for partial differentiation:

\[
\partial_\alpha \equiv \frac{\partial}{\partial x_\alpha}
\]

In this notation, and using the summation convention, the multivariate Taylor expansion looks particularly neat:

\[
f(x + h) = f(x) + h_\alpha \partial_\alpha f(x) + \frac{1}{2} h_i h_j \partial_i \partial_j f(x) + \frac{1}{3!} h_i h_j h_k \partial_i \partial_j \partial_k f(x) + \ldots
\]

(summation convention) The vectors \( h \) have been moved to the left-hand-side of all derivative operators to avoid any possible confusion about which quantity is differentiated.
7.5.4. Example: Derivative of $1/r$. The first order partial derivative of
\[ \frac{1}{r} \equiv \frac{1}{|r|} = \frac{1}{\sqrt{x^2 + y^2 + z^2}} \]
with respect to $x$ is:
\[ \frac{\partial}{\partial x} \left( \frac{1}{r} \right) = \frac{-1}{2} \frac{(2x)}{(x^2 + y^2 + z^2)^{3/2}} = -\frac{x}{r^3}. \]
Similar expressions are found for differentiation with respect to $y$ or $z$. Thus, for any component $\alpha = x, y, z$ we have:
\[ \frac{\partial}{\partial r_\alpha} \left( \frac{1}{r} \right) = \frac{\partial_\alpha}{r} \left( \frac{1}{r} \right) = -\frac{r_\alpha}{r^3} \]
where $r_\alpha$ can be any of $x, y, z$ depending on the choice of index $\alpha = x, y, z$.

**Problem 73.** (a) Is the genetic algorithm a “local” or “global” optimization algorithm? Explain.
(b) What element of this algorithm gives it its “local” or “global” nature?
(c) Discuss what is meant by local / global optimization, and give examples of algorithms (genetic or not) which achieve local vs. global optimization.

**Problem 74.** Solve analytically the following minimization problem for the unknown vector $g$:
\[ \arg \min_g \|Ug - k\|^2 \]
where $g$ is a column vector with $P$ entries whose values are unknown quantities that you are to solve for, $k$ is a column vector with $N$ rows. $U$ is a $N \times P$ matrix ($N$ rows, $P$ columns). The minimization is to be carried out by searching over all possible vector $g$ with finite entries.
(a) Find the exact solution to the minimization problem.
(b) Explain how this minimization problem can be used to fit experimental data to a model (linear or non-linear model). Explain what would be the roles of $U$, $g$ and $k$ in this context.

**Problem 75.** Solve the following equation giving the unknown $x$ number of moles of a substance needed in a reaction:
\[ 4 \sin(2x) + 5 \ln(2x^2) - 1000 + x^2 \exp(5x) = 0. \]
(a) Derive an algorithm and write a working computer program or use a spreadsheet or calculator to obtain a correct value for $x$. Explain how to do it.
(b) Find the value $x$ such that the left hand side of the equation is a minimum. You found a local minimum. Explain how you found a local minimum. Is it possible to find a global minimum?
(c) Find a minimum of the function \( f(x, y) = x^2 \cos(2x) \cos(2y) \) near the point \((x, y) = (100, 100)\). Explain all the details of how you proceeded to find the minimum.

(d) Fit data points \((95, 85), (85, 95), (80, 70), (70, 65), (60, 70)\) (data given in the form \((x_i, y_i)\)) to a model \( y(x) = A + Bx \) and find the values for \(A\) and \(B\). Show all your work.

**Solution.**

(a) Set \( f(x) = 4 \sin(2x) + 5 \ln(2x^2) - 1000 + x^2 \exp(5x) \) and use Newton Raphson \( x_{n+1} = x_n - f(x_n)/f'(x_n) \) using some initial guess \( x_0 \).

(b) \( f(x) = 4 \sin(2x) + 5 \ln(2x^2) - 1000 + x^2 \exp(5x) \) and use Newton Raphson applied to the first derivative, \( x_{n+1} = x_n - f'(x_n)/f''(x_n) \), and a suitable initial guess for \( x_0 \). You can find infinitely many local minima because of the \( \sin(2x) \) term. The function does not have a global minimum due to the singularity at the origin, \( \lim_{x \to 0} \ln(2x^2) = -\infty \).

(c) Use any of the gradient-based search algorithms (steepest descent, Newton, Gauss-Newton or Levenberg-Marquardt. Set initial conditions to \((100, 100)\).

(d) Use the formulae, \( A = \frac{\sum x_i^2 \sum y_i - \sum x_i \sum x_i y_i}{\Delta}, B = \frac{n \sum x_i y_i - \sum x_i \sum y_i}{\Delta}, \Delta = n \sum x_i^2 - (\sum x_i)^2 \), and plug in the data provided to compute \( A \) and \( B \).

I get the results: \( A = 26.768 \) and \( B = 0.644 \), i.e. the model is \( y(x) = 26.768 + 0.644x \).
Chapter 8

Errors in the Fitted Parameters during Nonlinear Fitting

Given that nonlinear fitting methods are based on computer algorithms, how can we obtain estimates of the error in the fitting parameters? We no longer have the option of deriving analytical formulas. It turns out that the covariance matrix provides estimates of the errors in the fitting parameters. In this lecture we show that $\text{cov}(\hat{\beta}_s, \hat{\beta}_s) = 2(H_s)^{-1}$, where $H_s$ is the Hessian matrix.

8.1. Linear least squares

In linear least squares, we assume a model of the form

$$\vec{y} = \underbrace{A\hat{\beta}}_{\text{model}} + \vec{\epsilon}$$

where $\vec{y} = (y_1, \ldots, y_n)^T$ is the column vector of measured data points. There are $n$ data points collected. $A$ is called the design matrix and has dimensions $n \times p$. The model parameters are stored in the vector $\hat{\beta} = (\beta_1, \ldots, \beta_p)^T$. Finally, the errors in each measurement are stored in a vector $\vec{\epsilon} = (\epsilon_1, \ldots, \epsilon_n)^T$.

The errors $\epsilon_i$ are random variables assumed to be independent and identically distributed (iid rv) according to a normal law

$$\epsilon_i \sim N(0, \sigma^2), \quad i = 1, \ldots, n.$$
In general \( n \gg p \), since we want a model with much fewer parameters than there are data points collected to construct it. In a sense, data fitting amounts to doing data compression. Some effort must be devoted to finding an adequate model that accurately describes the key features of the problem with as few parameters as possible. In components form, the above equation can be written as

\[
y_i = \sum_{j=1}^{p} a_{ij} \beta_j + \epsilon_i
\]

where \( a_{ij} \) are the elements of the design matrix \( A \).

Let us look at some examples of design matrix \( A \). This matrix is found in the expression \( \vec{y} = A \vec{\beta} + \vec{\epsilon} \). Consider the model

\[
y_i = A + B x_i + C x_i^2 + \epsilon_i, \quad i = 1, \ldots, n
\]

We find, by inspection

\[
A = \begin{bmatrix}
1 & x_1 & x_1^2 \\
1 & x_2 & x_2^2 \\
\vdots & \vdots & \vdots \\
1 & x_n & x_n^2 \\
\end{bmatrix}, \quad \vec{\beta} = \begin{bmatrix}
A \\
B \\
C \\
\end{bmatrix}, \quad \vec{y} = \begin{bmatrix}
y_1 \\
\vdots \\
y_n \\
\end{bmatrix}, \quad \vec{\epsilon} = \begin{bmatrix}
\epsilon_1 \\
\vdots \\
\epsilon_n \\
\end{bmatrix}
\]

Another example is

\[
y_i = A + B \log x_i + C \cos^2(x_i) + \epsilon_i
\]

We find, by inspection that

\[
A = \begin{bmatrix}
1 & \log x_1 & \cos^2(x_1) \\
1 & \log x_2 & \cos^2(x_2) \\
\vdots & \vdots & \vdots \\
1 & \log x_n & \cos^2(x_n) \\
\end{bmatrix}
\]

8.2. MLE

Recall that with least squares, the fitting parameters are obtained by applying the principle of maximum likelihood, which consists of solving for the model parameters \( \vec{\beta} \) for which the probability density of the observed deviations (data minus model) is a maximum. Maximizing \( L \) is equivalent to maximizing its logarithm, \( l \):

\[
l = -\frac{\chi^2}{2} = -\frac{1}{2} \sum_{i=1}^{n} \frac{(y_i - y(x_i; \vec{\beta}))^2}{\sigma_i^2} = \frac{1}{2} \sum_{i=1}^{n} \frac{\epsilon_i^2}{\sigma_i^2}
\]

Here, our assumption will be that \( \epsilon_i \) are Gaussian iid rv. Therefore, the joint probability distribution of all \( (\epsilon_1, \ldots, \epsilon_n) \) is a product of individual
distributions for each of the $\epsilon_i$, each of which is Gaussian with mean 0 and standard deviation $\sigma$.

In this special case, the “likelihood function” is

$$L(\beta, \sigma^2) \equiv p(\vec{y}|\vec{\beta}, \sigma^2) = \prod_{i=1}^{n} \exp\left\{ -\frac{1}{2} \left( y_i - \sum_{j=1}^{p} a_{ij} \beta_j \right)^2 / \sigma^2 \right\} / (\sigma^2 (2\pi)^{n/2})$$

$L$ is the probability density for observing the set of deviations $\{\epsilon_i\}$ or, equivalently, the set of measurements $\vec{y}$ given a model and its parameters $y(x_i; \vec{\beta})$.

Taking the natural log of this expression yields the “log likelihood”

$$l \equiv \log L = \text{const} - n \log \sigma - \frac{1}{2} \left[ \sum_{i=1}^{n} \left( y_i - \sum_{j=1}^{p} a_{ij} \beta_j \right)^2 / \sigma^2 \right]$$

where $\text{const} = -(n/2) \log(2\pi)$. This term will be ignored since it will not be needed.

The principle of “maximum likelihood” tells us we should compute the extremum of this function with respect to its parameters. The parameters are $\vec{\beta}$ and $\sigma$.

### 8.2.1. MLE of the Model Parameters $\vec{\beta}$

Differentiating with respect to $\vec{\beta}$, i.e. $\partial l / \partial \beta_r = 0$, gives the following equation:

$$\sum_{i=1}^{n} a_{ir} \left( y_i - \sum_{j=1}^{p} a_{ij} \beta_j \right) = 0, \quad (r = 1, \ldots, p)$$

or,

$$\sum_{j=1}^{p} \left( \sum_{i=1}^{n} a_{ir} a_{ij} \right) \beta_j = \sum_{i=1}^{n} a_{ir} y_i.$$

This can be expressed in a more compact notation if we write $C = (c_{ij})$ for the $ij$-th element of the $p \times p$ matrix $C \equiv A^T A$. We note that this matrix is symmetric ($C^T = C$) since $C^T = (A^T A)^T = A^T A = C$. This allows us to write the bracketed term on left hand side as

$$c_{ij} = \sum_{k=1}^{n} (A^T)_{ik} A_{kj} = \sum_{k=1}^{n} a_{ki} a_{kj}$$

so the left hand side becomes

$$\sum_{j=1}^{p} c_{rj} \beta_j = \sum_{i=1}^{n} a_{ir} y_i = \sum_{i=1}^{n} (A^T)_{rI} y_i$$
or
\[
(C\vec{\beta})_r = (A^T \vec{y})_r, \quad r = 1, \ldots, p
\]
Thus, we have obtained the so-called “normal equations”
\[
C\vec{\beta} = A^T \vec{y} \quad C = A^T A
\]
The solution to these normal equations is obtained by multiplying both sides of the equation on the left by the inverse of C, which is simply $C^T$ (since C is a symmetric matrix). Since the vector $\tilde{\beta}$ has been obtaining from the extremum condition $\partial l/\partial \vec{\beta} = 0$ we denote this particular vector as $\hat{\beta}$:
\[
\hat{\beta} = C^{-1}A^T \vec{y} = (A^T A)^{-1} A^T \vec{y}
\]
provided that the symmetric matrix $C = A^T A$ is invertible. We note that the normal equations are linear in $\vec{y}$. They are also linear in $\vec{\beta}$, meaning that we can easily solve for $\beta$. In the case of nonlinear models, we will see that the normal equations are generally not linear functions of the parameters.
Finally, to summarize what we have done, we note that maximizing the likelihood function $L$ is entirely equivalent to minimizing the sum of square errors with respect to the choice of $\vec{\beta}$
\[
\sum_{i=1}^{n} \left( y_i - \sum_{j=1}^{p} a_{ij}\beta_j \right)^2 = (\vec{y} - A\vec{\beta})^T (\vec{y} - A\vec{\beta}) = \vec{\epsilon}^T \vec{\epsilon}.
\]
The solution to this problem is given by the solution to the normal equations. The term $\vec{\epsilon}^T \vec{\epsilon}$ is an inner product of $\vec{\epsilon}$ with itself. This is often expressed in terms of the norm (length) of the vector $\vec{\epsilon}$ as follows: $\vec{\epsilon}^T \vec{\epsilon} = ||\vec{\epsilon}||^2$, where $||\vec{\epsilon}||$ is the length of $\vec{\epsilon}$.
To illustrate the use of the normal equations in solving linear least squares problems, let us look at the example of a linear model ($y_i = A + Bx_i$). This equation must be written in the form $\vec{y}(\vec{x}; \vec{\beta}) = A\vec{\beta}$ where $A$ is the design matrix. By inspection we see that
\[
\vec{y}(\vec{x}; \vec{\beta}) = A\vec{\beta} = \begin{bmatrix} x_1 & 1 \\ x_2 & 1 \\ \vdots \\ x_n & 1 \end{bmatrix} \begin{bmatrix} B \\ A \end{bmatrix} = \begin{bmatrix} y_1 \\ y_2 \\ \vdots \\ y_n \end{bmatrix}
\]
The experimentally measured data points $\vec{y}$ are expressed as the sum of the model $\vec{y}(\vec{x}; \vec{\beta}) = A\vec{\beta}$ plus the random errors $\vec{\epsilon}$
\[
\vec{y} = A\vec{\beta} + \vec{\epsilon}
\]
where
\[
\vec{y} = \begin{bmatrix} y_1 \\ y_2 \\ \vdots \\ y_n \end{bmatrix}, \quad \vec{\beta} = \begin{bmatrix} B \\ A \end{bmatrix}, \quad \vec{\epsilon} = \begin{bmatrix} \epsilon_1 \\ \epsilon_2 \\ \vdots \\ \epsilon_n \end{bmatrix}
\]

This is nothing more than a restatement of the fact that \( \vec{\epsilon} \) is simply the deviation \( \vec{y} - A \vec{\beta} \) of experimentally measured data \( \vec{y} \) from the model \( A \vec{\beta} \). That is, \( \vec{\epsilon} = \vec{y} - A \vec{\beta} \). We recall that \( \chi^2 \) involves a summation over such deviations.

We now compute the product \( C^{-1} A^T \vec{y} \), where \( C = A^T A \). First, the products \( C = A^T A \) and \( A^T \vec{y} \) are, respectively
\[
C = A^T A = \begin{bmatrix} x_1 & x_2 & \cdots & x_n \\ 1 & 1 & \cdots & 1 \end{bmatrix} \begin{bmatrix} x_1 \\ x_2 \\ \vdots \\ x_n \\ 1 \end{bmatrix} = \begin{bmatrix} \sum_i x_i^2 & \sum_i x_i \\ \sum_i x_i & n \end{bmatrix},
\]
and
\[
A^T \vec{y} = \begin{bmatrix} \sum_i x_i y_i \end{bmatrix}.
\]

Next, the inverse \( C^{-1} \) of this 2x2 matrix is computed from the usual formula
\[
\begin{pmatrix} a & b \\ c & d \end{pmatrix}^{-1} = \frac{1}{ad - bc} \begin{pmatrix} d & -b \\ -c & a \end{pmatrix}.
\]

You can easily check by direct multiplication that this matrix meets the conditions required to be the inverse: \( C^{-1} C = C C^{-1} = (1 0) \).

In our case, the inverse of \( C = \begin{bmatrix} \sum_i x_i^2 & \sum_i x_i \\ \sum_i x_i & n \end{bmatrix} \) is
\[
C^{-1} = \frac{1}{n \sum_i x_i^2 - \sum_i x_i \sum_j x_j} \begin{bmatrix} n & -\sum_i x_i \\ -\sum_i x_i & \sum_i x_i^2 \end{bmatrix}.
\]

Therefore, \( \hat{\beta} = \begin{bmatrix} \hat{B} \\ \hat{A} \end{bmatrix} \) is equal to
\[
C^{-1} A^T \vec{y} = \frac{1}{n \sum_i x_i^2 - \sum_i x_i \sum_j x_j} \begin{bmatrix} n \sum_i x_i y_i - \sum_i x_i \sum_j y_j \\ -\sum_i x_i \sum_j x_j y_j + \sum_i x_i^2 \sum_j y_j \end{bmatrix}.
\]

This result is nothing new: this is exactly the same result we have derived in Lecture 4 by applying the maximum likelihood principle to the minimization of \( \chi^2 \). Here, we simply verified that our normal equation is in agreement with the result already derived previously.
8.3. MLE of the Parameter $\sigma$

There is yet another parameter that can be estimated via maximum likelihood, $\sigma$. While $\sigma$ does not appear at first glance to be a "model" parameter in the sense of the model expressed as $\tilde{y}(\tilde{x}; \tilde{\beta}) = A(\tilde{x})\tilde{\beta}$, we initially stated the assumption that our deviations $\tilde{e} = \tilde{y} - \tilde{y}(\tilde{x}; \tilde{\beta})$ were distributed as Gaussians with mean 0 and standard deviation $\sigma$. In this sense, $\sigma$ is an unknown parameter that can be solved for in terms of the data and the model.

Differentiation of the log likelihood

$$l = \log L = \text{const} - n \log \sigma - \frac{1}{2} \left[ \sum_{i=1}^{n} \left( y_i - \sum_{j=1}^{p} a_{ij} \hat{\beta}_j \right)^2 \right] / \sigma^2,$$

with respect to the parameter $\sigma$ gives

$$\frac{\partial l}{\partial \sigma} = 0 \quad \Rightarrow \quad -\frac{n}{\sigma} + \frac{1}{\sigma^3} \sum_{i=1}^{n} \left( y_i - \sum_{j=1}^{p} a_{ij} \hat{\beta}_j \right)^2 = 0.$$

At the extremum, we will denote the value of $\tilde{\beta}$ by $\hat{\beta}$. Similarly, let us denote the value of $\sigma$ as $\hat{\sigma}$ at the extremum. The quantities $\hat{\beta}$ and $\hat{\sigma}$ are called maximum likelihood estimators. Solving for $\sigma^2$ we get, for the maximum likelihood estimator:

$$\hat{\sigma}^2 = \frac{1}{n} \sum_{i=1}^{n} \left( y_i - \sum_{j=1}^{p} a_{ij} \hat{\beta}_j \right)^2.$$

8.4. The Covariance Matrix

An important concept that will be used later in this course is the covariance matrix. We recall near the end of Lecture 2 we had defined the covariance of two random variables $X$ and $Y$:

$$\text{cov}(X, Y) = (X - \overline{X})(Y - \overline{Y}) = \overline{XY} - \overline{X} \cdot \overline{Y}.$$

Covariance is important because in the case $Y = X$ it reports the variance whereas for $X \neq Y$ it tell us to what extent the random variable $X$ is correlated to $Y$. The covariance matrix is used when $X$ is a vector in which we have collected all random variables of interest, i.e. $\tilde{X} = (X, Y, \ldots, Z)^T$. It is defined as the matrix whose elements are $\text{cov}(X_i, X_j)$. It can be computed most easily using the matrix form

$$\text{cov}(\tilde{X}, \tilde{X}) = (\tilde{X} - \overline{\tilde{X}})(\tilde{X} - \overline{\tilde{X}})^T.$$
Here, we are interested in the covariance matrix of the fitted model parameters $\text{cov}(\hat{\beta}, \hat{\beta})$. The diagonal elements are the variances (errors squared) of each model parameter. The off-diagonal elements are the covariances. The matrix elements are important because we are interested in knowing from the fitting procedure what are the uncertainties in each fitted parameter and the extent to which model parameters are redundant (correlated). Using the definition $\hat{\beta} = C^{-1}A^T \vec{y}$ we find:

$$\text{cov}(\hat{\beta}, \hat{\beta}) = C^{-1}A^T (\vec{y} - \bar{y})(C^{-1}A^T (\vec{y} - \bar{y}))^T = C^{-1}A^T (\vec{y} - \bar{y})(\vec{y} - \bar{y})^T A(C^{-1})^T.$$

We recognize the term $(\vec{y} - \bar{y})(\vec{y} - \bar{y})^T$ as a variance. Since the elements of $\vec{y}$ are statistically independent with variance $\sigma^2$ (by assumption), the off-diagonal elements are zero and this variance evaluates to

$$\begin{bmatrix}
\sigma^2 & 0 & \ldots & 0 \\
0 & \sigma^2 & \ldots & 0 \\
\vdots & \vdots & \ddots & \vdots \\
0 & 0 & \ldots & \sigma^2
\end{bmatrix} = \sigma^2 \mathbf{1},$$

where $\mathbf{1}$ is the unit matrix. In the second line we have used the property $(AB)^T = B^T A^T$ to write the term $(C^{-1}A^T (\vec{y} - \bar{y}))^T$ as

$$(C^{-1}A^T)^T (\vec{y} - \bar{y}) = (\vec{y} - \bar{y})^T A(C^{-1})^T.$$

Next, we use the fact that inverse of a matrix and transpose operations are interchangeable, i.e. $(C^T)^{-1} = (C^{-1})^T$. But since $C = C^T$ (symmetric matrix) we have:

$$\text{cov}(\hat{\beta}, \hat{\beta}) = C^{-1}A^T \sigma^2 \mathbf{1} C^{-1} = \sigma^2 C^{-1} A C^{-1} = \sigma^2 C^{-1}.$$

We will see later that the covariance matrix is related to the inverse of the Hessian matrix. This is useful because we have seen in previous lectures that the Hessian matrix could be estimated from the Jacobian matrix during the course of a data fitting procedure (c.f. Gauss-Newton method).

\footnote{Recall that $\vec{y} = A\vec{\beta} + \vec{\epsilon}$ is a sum of a deterministic (non-random) quantity $A\vec{\beta}$ plus a random quantity $\vec{\epsilon}$ which we have assumed the elements of which to be iid rvs with $N(0, \sigma^2)$. Thus, $\text{var}(\vec{y}) = \text{var}(\vec{\epsilon}) = \sigma^2$.}
This formula $\text{cov}(\hat{\beta}, \hat{\beta}) = \sigma^2 \mathbf{C}^{-1}$ we have derived is simply a tool that enables us to compute the covariance matrix $\text{cov}(\hat{\beta}, \hat{\beta})$ from the conditions of the experiment. The covariance matrix is important because its diagonal elements are the variances in the fitted parameters. Let us look at a simple example: that of the linear model $y_i = A + Bx_i$. Then, $\vec{\beta} = (B, A)^T$ and therefore

\[
\text{cov}(\hat{\beta}, \hat{\beta}) = \begin{bmatrix}
(B - \bar{B})(B - \bar{B}) & (B - \bar{B})(A - \bar{A}) \\
(A - \bar{A})(B - \bar{B}) & (A - \bar{A})(A - \bar{A})
\end{bmatrix}
= \begin{bmatrix}
\text{var}(B) & \text{cov}(A, B) \\
\text{cov}(B, A) & \text{var}(A)
\end{bmatrix}.
\]

We have also seen that $\mathbf{C}^{-1} = \frac{1}{n \sum_i x_i^2 - \sum_i x_i \sum_j x_j} \left[ n \sum_i x_i - \sum_j x_j^2 \right]$. Thus, we have an explicit expression for the formula $\text{cov}(\hat{\beta}, \hat{\beta}) = \sigma^2 \mathbf{C}^{-1}$ and can use it to compute the errors in the fitted parameters $A$ and $B$ from the experimental conditions (design matrix). We note that the \{\{y_i\}\} dependence of the errors and covariances originates from the factor $\sigma^2$.

### 8.5. Nonlinear Least Squares

Suppose that we have $n$ observations $(\vec{x}_i, y_i)$, $i = 1, \ldots, n$ and a nonlinear model

\[y_i = f(\vec{x}_i; \vec{\beta}) + \epsilon_i, \quad i = 1, \ldots, n\]

or

\[\vec{y} = f(\vec{x}; \vec{\beta}) + \vec{\epsilon},\]

where $f(\vec{x}; \vec{\beta})$ is a nonlinear function of $\vec{\beta}$. We assume that the elements of $\vec{\epsilon}$ are independent identically distributed random variables with mean 0 and variance $\sigma^2$. Furthermore, we assume they are normally distributed. Thus, the estimates $\hat{\beta}$ will also be known as maximum likelihood estimators.

Given this nonlinear model, let us obtain the least squares estimates of $\vec{\beta}$ and $\sigma$, which we denote $\hat{\beta}$ and $\hat{\sigma}$ respectively, starting from the likelihood function

\[L(\vec{\beta}, \sigma^2) \equiv p(\vec{y} | \vec{\beta}, \sigma^2) = \frac{1}{(2\pi\sigma^2)^{n/2}} \exp\left(-\frac{1}{2} \sum_{i=1}^{n} \frac{[y_i - f(\vec{x}_i; \vec{\beta})]^2}{\sigma^2}\right)\]
The log likelihood is (ignoring constants)
\[
l(\bar{\beta}, \sigma^2) = -\frac{n}{2} \log \sigma^2 - \frac{1}{2\sigma^2} \sum_{i=1}^{n} [y_i - f(\bar{x}_i; \bar{\beta})]^2
\]

### 8.5.1. MLE \(\hat{\sigma}^2\)

The condition \(\partial l / \partial \sigma^2 = 0\) gives the maximum likelihood estimator:
\[
\hat{\sigma}^2 = \frac{1}{n} \sum_{i=1}^{n} [y_i - f(\bar{x}_i; \bar{\beta})]^2
\]
which is indeed a "maximum" likelihood (for a given \(\bar{\beta}\)) as the second derivative is negative\(^2\):
\[
\frac{\partial^2 l}{\partial (\sigma^2)^2} = \frac{\partial}{\partial \sigma^2} \left[ -\frac{n}{2\sigma^2} + \sum_{i=1}^{n} \frac{[y_i - f(\bar{x}_i; \bar{\beta})]^2}{2(\sigma^2)^2} \right] = \frac{n}{2(\sigma^2)^2} - \frac{1}{(\sigma^2)^3} \sum_{i=1}^{n} [y_i - f(\bar{x}_i; \bar{\beta})]^2
\]
in which we substitute the value of \(\sigma^2\), namely \(\bar{\epsilon}^T \bar{\epsilon} = \sum_{i=1}^{n} [y_i - f(\bar{x}_i; \bar{\beta})]^2\) to get
\[
\frac{\partial^2 l}{\partial (\sigma^2)^2} = \frac{n^3}{2(\bar{\epsilon}^T \bar{\epsilon})^2} - \frac{n^3}{(\bar{\epsilon}^T \bar{\epsilon})^3} \cdot (\bar{\epsilon}^T \bar{\epsilon}) = \frac{n^3}{(\bar{\epsilon}^T \bar{\epsilon})^2} (\frac{1}{2} - 1) < 0
\]
hence \(\hat{\beta}\) and \(\hat{\sigma}^2\) maximize \(l(\bar{\beta}; \sigma^2)\). The maximum value of \(p(\bar{y}|\bar{\beta}, \sigma^2)\) is
\[
p(\bar{y}|\hat{\beta}, \hat{\sigma}^2) = \frac{1}{(2\pi\hat{\sigma}^2)^{n/2}} e^{-n/2}.
\]

### 8.5.2. MLE \(\hat{\beta}\)

The estimator \(\hat{\beta}\) satisfies the condition \(\partial \bar{\epsilon}^T \bar{\epsilon} / \partial \beta_r |_{\bar{\beta}} = 0\), \(r = 1, 2, \ldots, p\). We now write \(f_i(\bar{\beta}) = f(\bar{x}_i; \bar{\beta})\) as shorthand notation. Then,
\[
\bar{f}(\bar{\beta}) = (f_1(\bar{\beta}), \ldots, f_n(\bar{\beta}))^T.
\]
Since \(f\) is a vector, its gradient is a matrix:
\[
\mathbf{F}(\bar{\beta}) \equiv \nabla \bar{f}(\bar{\beta}) = \frac{\partial \bar{f}(\bar{\beta})}{\partial \bar{\beta}} = \begin{bmatrix}
\frac{\partial f_1}{\partial \beta_1} & \cdots & \frac{\partial f_1}{\partial \beta_p} \\
\frac{\partial f_2}{\partial \beta_1} & \cdots & \frac{\partial f_2}{\partial \beta_p} \\
\vdots & \ddots & \vdots \\
\frac{\partial f_n}{\partial \beta_1} & \cdots & \frac{\partial f_n}{\partial \beta_p}
\end{bmatrix}.
\]
We now use the shorthand notation \(\mathbf{F} = \mathbf{F}(\bar{\beta})\) and \(\hat{\mathbf{F}} = \mathbf{F}(\hat{\beta})\). The sum of squares that must be minimized is
\[
\bar{\epsilon}^T \bar{\epsilon}(\bar{\beta}) = [\bar{y} - \bar{f}(\bar{\beta})]^T [\bar{y} - \bar{f}(\bar{\beta})] = ||\bar{y} - \bar{f}(\bar{\beta})||^2.
\]
\(^2\)Another way to see this, since \(\bar{\epsilon}^T \bar{\epsilon}(\bar{\beta}) \geq \bar{\epsilon}^T \bar{\epsilon}(\hat{\beta})\), we have \(l(\hat{\beta}) - l(\bar{\beta}) = -(n/2) \log \hat{\sigma}^2 - (n/2) - l(\hat{\beta}) \geq -(n/2) \log \hat{\sigma}^2 \sigma^2 - (n/2) + (1/2) \bar{\epsilon}^T \bar{\epsilon}(\hat{\beta}) / \sigma^2 = -(n/2) \log \hat{\sigma}^2 / \sigma^2 + 1 - \hat{\sigma}^2 / \sigma^2 \geq 0\) as \(\log x \leq x - 1\) for \(x \geq 0\). We have denoted \(\bar{\epsilon}^T \bar{\epsilon}(\hat{\beta})\) for \(\bar{\epsilon}^T \bar{\epsilon}\) evaluated at the point \(\hat{\beta}\).
where $\hat{\epsilon}^T \hat{\epsilon} (\hat{\beta})$ denotes $\hat{\epsilon}^T \hat{\epsilon}$ evaluated at the point $\hat{\beta}$. Setting $\frac{\partial \hat{\epsilon}^T \hat{\epsilon} (\hat{\beta})}{\partial \beta_r} |_{\hat{\beta}} = 0$ leads to the result
\[
\sum_i \{ y_i - f_i(\hat{\beta}) \} \frac{\partial f_i(\hat{\beta})}{\partial \beta_r} \bigg|_{\hat{\beta}} = 0 \Rightarrow 0 = \hat{F}^T \{ \hat{y} - \hat{f}(\hat{\beta}) \}
\]

We get the normal equations for this nonlinear model
\[
0 = \hat{F}^T \cdot \hat{\epsilon}.
\]

This is the equation that must be solved for nonlinear models. We note that the model parameters are contained implicitly in the matrix $F$ as partial derivatives. Thus, if the model $\hat{f}$ is nonlinear, the matrix elements of $F$ could also be nonlinear functions of the parameters $\hat{\beta}$.

In general, these normal equations cannot be solved analytically and iterative methods of the type covered previously will be necessary.

In order to better understand how to use the normal equations we illustrate this concept by looking at examples.

Consider the linear model
\[
y_i \underbrace{\text{data}}_{y} = A + Bx_i + \epsilon_i \underbrace{\text{model}}_{\epsilon}
\]

The normal equations are $0 = \hat{F}^T \cdot \hat{\epsilon}$. The vector $\hat{\epsilon}$ is simply the vector of deviations between experimental data $\hat{y}$ and the model $\hat{y}(x; \hat{\beta})$:
\[
\hat{\epsilon} = \hat{y} - \hat{y}(\hat{x}; \hat{\beta}) = \hat{y} - A - B\hat{x} = \begin{bmatrix} y_1 - A - Bx_1 \\ y_2 - A - Bx_2 \\ \vdots \\ y_n - A - Bx_n \end{bmatrix}
\]
or, in component form, $\epsilon_j = y_i - A - Bx_i$. Next, we need the matrix $\hat{F}^T$,

which is computed from $F(\hat{\beta})^T = \begin{bmatrix} \frac{\partial f_1}{\partial \beta_1} & \cdots & \frac{\partial f_n}{\partial \beta_1} \\ \frac{\partial f_1}{\partial \beta_2} & \cdots & \frac{\partial f_n}{\partial \beta_2} \\ \vdots & \cdots & \vdots \\ \frac{\partial f_1}{\partial \beta_p} & \cdots & \frac{\partial f_n}{\partial \beta_p} \end{bmatrix}$, where $f_i = A + Bx_i$ and $\hat{\beta} = (A, B)^T$, but with two parameters ($p = 2$)
\[
F(\hat{\beta})^T = \begin{bmatrix} \frac{\partial f_1}{\partial \beta_1} & \cdots & \frac{\partial f_n}{\partial \beta_1} \\ \frac{\partial f_1}{\partial \beta_2} & \cdots & \frac{\partial f_n}{\partial \beta_2} \end{bmatrix} = \begin{bmatrix} 1 & \cdots & 1 \\ x_1 & \cdots & x_n \end{bmatrix}
\]
The condition \( \hat{\mathbf{F}}^T \cdot \hat{\epsilon} = 0 \) then reads
\[
\sum_i (y_i - A - B x_i) = 0
\]
\[
\sum_i (y_i - A - B x_i) x_i = 0
\]
The two linear equations and two unknowns can be solved for \( A, B \) (taking care of noting that the second term in the first equation is \( \sum_i A = n A \)):
\[
B = \frac{n \sum_i y_i x_i - \sum_i x_i \sum_j y_j}{n \sum_i x_i^2 - \sum_i x_i \sum_j x_j}
\]
\[
A = \frac{\sum_i x_i^2 \sum_j y_j - \sum_i y_i x_i \sum_j x_j}{n \sum_i x_i^2 - \sum_i x_i \sum_j x_j}
\]
This is identical to the solution derived earlier using the normal equations for the linear case.

Consider the nonlinear model
\[
y_i = \alpha x_i^\beta + \epsilon_i \quad (i = 1, \ldots, n)
\]
This model can describe, for example, the dependence of mass of an object \( (y) \) on a side length \( (x) \) in \( \beta \) dimensions. The normal equations are:
\[
\sum_i (y_i - \alpha x_i^\beta) x_i^\beta = 0
\]
\[
\sum_i (y_i - \alpha x_i^\beta) \alpha x_i^\beta \log x_i = 0
\]
These equations do not admit analytical solutions for \( \alpha \) and \( \beta \).

8.6. Linearizing a nonlinear model

In the nonlinear case, the minimization of \( \| \hat{y} - \tilde{f}(\hat{\beta}) \|^2 \) with respect to \( \hat{\beta} \) yielded the normal equation \( 0 = \hat{\mathbf{F}}^T \cdot \hat{\epsilon} \), where \( \mathbf{F} = \partial \tilde{f} / \partial \hat{\beta} \) is generally nonlinear in \( \hat{\beta} \) and \( \hat{\epsilon} = \hat{y} - \tilde{f}(\hat{\beta}) \) is also nonlinear in \( \beta \). These “normal equations”, although formally correct, generally cannot be solved algebraically. In many cases, we resort to linearizing the nonlinear model, as we have done with the Newton-Raphson method. This approach does not yield the solution in a single step, but instead allows us to get progressively closer after each iteration of an update rule.

In order to do this, we first recall the following tool that we will use later. We have previously found that the minimization of \( \| \tilde{\epsilon} \|^2 = \| \hat{y} - \mathbf{A} \hat{\beta} \|^2 \) has the following solution
\[
\hat{\beta} = (\mathbf{A}^T \mathbf{A})^{-1} \mathbf{A}^T \hat{y}.
\]
This was obtained by minimizing \( \| \tilde{\epsilon} \|^2 \) with respect to \( \hat{\beta} \), i.e. \( \partial \| \tilde{\epsilon} \|^2 / \partial \hat{\beta} = 0 \). This solution is not specific to data fitting, but is a general result from calculus.
8. Errors in the Fitted Parameters during Nonlinear Fitting

Taylor expanding $\tilde{f}(\hat{\beta})$ near $\hat{\beta}$ to first order gives $\tilde{f}(\hat{\beta}) \approx \tilde{f}(\hat{\beta}) + F\Delta\hat{\beta}$ where $\Delta\hat{\beta} = \hat{\beta} - \hat{\beta}$. Thus,

$$\|\tilde{y} - \tilde{f}(\hat{\beta})\|^2 \approx \|\tilde{y} - \tilde{f}(\hat{\beta}) - F\Delta\hat{\beta}\|^2.$$ 

Substitution of $\hat{\epsilon} = \tilde{y} - \tilde{f}(\hat{\beta})$ for $\tilde{f}(\hat{\beta})$ yields

$$\|\hat{\epsilon} - F\Delta\hat{\beta}\|^2.$$ 

Invoking the above theorem from calculus, we find that the solution of this linearized problem is given by

$$\Delta\hat{\beta} = (F^TF)^{-1}F^T\hat{\epsilon}.$$ 

Incidentally, this condition

$$\Delta\hat{\beta} = \hat{\beta} - \hat{\beta} = (F^TF)^{-1}F^T\hat{\epsilon} \quad (\star)$$ 

is exactly equivalent to the update rule we derived previously for the Gauss-Newton method

$$\hat{\beta}_{s+1} = \hat{\beta}_s - H_s^{-1}\nabla^2\chi^2(\hat{\beta}_s)$$ 

where the Hessian matrix $H_s$ is approximated by $2J^TJ$ with

$$J = J(\hat{\beta}_s) = \begin{bmatrix} -\frac{1}{\sigma_1}\frac{\partial y(x_1;\hat{\beta}_s)}{\partial \beta_1} & \cdots & -\frac{1}{\sigma_1}\frac{\partial y(x_1;\hat{\beta}_s)}{\partial \beta_p} \\ \vdots & \ddots & \vdots \\ -\frac{1}{\sigma_n}\frac{\partial y(x_n;\hat{\beta}_s)}{\partial \beta_1} & \cdots & -\frac{1}{\sigma_n}\frac{\partial y(x_n;\hat{\beta}_s)}{\partial \beta_p} \end{bmatrix}$$ 

and $\nabla^2\chi^2 = 2J^T\tilde{R}$ with $\tilde{R}$ is the column vector of residuals $\tilde{R}_i = (y_i - y(x_i;\hat{\beta}_s))/\sigma_i$. For the two rules to be equivalent we must take the errors to be identical $\sigma_i = \sigma$. In that case,

$$H_s^{-1}2J^T \begin{bmatrix} (y_1 - y(x_1;\hat{\beta}_s))/\sigma \\ \vdots \\ (y_n - y(x_n;\hat{\beta}_s))/\sigma \end{bmatrix} = (2J^TJ)^{-1}2J^T \begin{bmatrix} (y_1 - y(x_1;\hat{\beta}_s))/\sigma \\ \vdots \\ (y_n - y(x_n;\hat{\beta}_s))/\sigma \end{bmatrix}$$ 

and we can now write $J = -(1/\sigma)F$ to get

$$= -\sigma^2(2F^TF)^{-1}2(1/\sigma)F^T \begin{bmatrix} (y_1 - y(x_1;\hat{\beta}_s))/\sigma \\ \vdots \\ (y_n - y(x_n;\hat{\beta}_s))/\sigma \end{bmatrix} = -(F^TF)^{-1}F^T\hat{\epsilon}$$ 

which gives an update rule identical to $(\star)$

$$\hat{\beta}_{s+1} = \hat{\beta}_s + (F^TF)^{-1}F^T\hat{\epsilon}$$ 

8.7. Relationship between Hessian and Covariance Matrices

The covariance matrix, $\text{cov}(\hat{\beta},\hat{\beta})$, is important because it contains information about the statistics of the fitted parameters. Let us see how it can be
computed in the case of an iterative algorithm for nonlinear data fitting. In
the linear case, we found an explicit formula:
\[
cov(\vec{\beta}, \vec{\beta}) = \sigma^2 \mathbf{C}^{-1} = \sigma^2 (\mathbf{A}^T \mathbf{A})^{-1}.
\] (linear case)

In the nonlinear case, we can linearize such that the matrix \( \mathbf{A} \) becomes\( \mathbf{F} = \partial \vec{f} / \partial \vec{\beta} \), the gradient of \( \vec{f} \):
\[
cov(\vec{\beta}, \vec{\beta}) = \sigma^2 (\mathbf{F}^T \mathbf{F})^{-1}.
\] (nonlinear case)

Let us recall the definition of \( \chi^2 \)
\[
\chi^2 = \sum_i (y_i - f_i(\vec{\beta}))^2 = \left\| \vec{y} - \vec{f}(\vec{\beta}) \right\|^2 = \frac{1}{\sigma^2} \vec{e}^T \vec{e}
\]
and compute the Hessian matrix,
\[
(H)_{ij} = \frac{\partial^2 \chi^2}{\partial \beta_i \partial \beta_j} = \frac{\partial^2}{\partial \beta_i \partial \beta_j} \left( \frac{1}{\sigma^2} \left\| \vec{y} - \vec{f}(\vec{\beta}) \right\|^2 \right)
\]
\[
= \frac{\partial^2}{\partial \beta_i \partial \beta_j} \left( \vec{y} - \vec{f}(\vec{\beta}) - \mathbf{F} \Delta \vec{\beta} \right)^T \left( \vec{y} - \vec{f}(\vec{\beta}) - \mathbf{F} \Delta \vec{\beta} \right) \frac{1}{\sigma^2}
\]
\[
= \frac{\partial^2}{\partial \beta_i \partial \beta_j} \left( \vec{y} - \vec{f}(\vec{\beta}) \right)^T \left( \vec{y} - \vec{f}(\vec{\beta}) \right) + 2(-\mathbf{F} \Delta \vec{\beta})^T (\vec{y} - \vec{f}(\vec{\beta})) + (\mathbf{F} \Delta \vec{\beta})^T (\mathbf{F} \Delta \vec{\beta}) \frac{1}{\sigma^2}
\]

The first term, \( (\vec{y} - \vec{f}(\vec{\beta}))^T (\vec{y} - \vec{f}(\vec{\beta})) \), depends on \( \vec{\beta} \) but is independent of \( \hat{\beta} \) and its derivative with respect to \( \vec{\beta} \) is therefore zero. The second term, \( 2(-\mathbf{F} \Delta \vec{\beta})^T (\vec{y} - \vec{f}(\vec{\beta})) \), depends linearly on \( \beta \), and so its second derivative
with respect to \( \vec{\beta} \) vanishes. This leaves only the third term as non-zero. Using the facts that \( \Delta \vec{\beta} = \vec{\beta} - \hat{\vec{\beta}} \) and \( \mathbf{F} = \partial \vec{f}(\vec{\beta}) / \partial \vec{\beta} |_{\vec{\beta} = \hat{\vec{\beta}}} \) is independent of \( \hat{\vec{\beta}} \), we compute the derivative:
\[
(H)_{ij} = \frac{\partial^2}{\partial \beta_i \partial \beta_j} \left( \mathbf{F} \Delta \vec{\beta} \right)^T (\mathbf{F} \Delta \vec{\beta}) \frac{1}{\sigma^2} = \frac{1}{\sigma^2} \partial^2_{\beta_i \beta_j} \left( \mathbf{F}^T \mathbf{F} \right) \Delta \vec{\beta}
\]
\[
= \frac{1}{\sigma^2} \partial^2_{\beta_i \beta_j} \sum_{m,n} \beta_m (\mathbf{F}^T \mathbf{F})_{mn} \beta_n
\]
\[
= \frac{1}{\sigma^2} \partial_{\beta_i} \left( \sum_n (\mathbf{F}^T \mathbf{F})_{jn} \beta_n + \sum_m \beta_m (\mathbf{F}^T \mathbf{F})_{mj} \right)
\]
\[
= \frac{1}{\sigma^2} (\mathbf{F}^T \mathbf{F})_{ji} + (\mathbf{F}^T \mathbf{F})_{ij} = \frac{2}{\sigma^2} (\mathbf{F}^T \mathbf{F})_{ij},
\]
where in the second equality we have replaced \( \Delta \vec{\beta} \) by \( \vec{\beta} \) because the derivative operation with respect to \( \vec{\beta} \). This leads to:
\[
\mathbf{H} = (2/\sigma^2)(\mathbf{F}^T \mathbf{F}).
\]
8.7.1. Summary: relationship between Hessian and covariance matrix. Thus, we have shown that the covariance matrix is proportional to the inverse of the Hessian

\[ \text{cov}(\vec{\beta}_s, \vec{\beta}_s) = 2(H_s)^{-1} \]

(The factor of 2 would go away had we defined the Hessian as the derivative of the log likelihood function; can you show this?) Recall that the covariance matrix contains variances of the fitted parameters $\vec{\beta}$ along the diagonal and covariances of these parameters as off-diagonal elements:

\[
\text{cov}(\vec{\beta}, \vec{\beta}) = \begin{bmatrix}
\text{var}(\beta_1) & \text{cov}(\beta_1, \beta_2) & \ldots & \text{cov}(\beta_1, \beta_p) \\
\text{cov}(\beta_2, \beta_1) & \text{var}(\beta_2) & \ldots & \text{cov}(\beta_2, \beta_p) \\
\vdots & \vdots & \ddots & \vdots \\
\text{cov}(\beta_p, \beta_1) & \text{cov}(\beta_p, \beta_2) & \ldots & \text{var}(\beta_p)
\end{bmatrix}
\]

For example, if we require the error bars on the fitted parameters, we may extract the diagonal elements of the matrix $2(H_s)^{-1}$

\[
\text{diag}(\text{cov}(\vec{\beta}, \vec{\beta})) = \begin{bmatrix}
\text{var}(\beta_1) & 0 & \ldots & 0 \\
0 & \text{var}(\beta_2) & \ldots & 0 \\
\vdots & \vdots & \ddots & \vdots \\
0 & 0 & \ldots & \text{var}(\beta_p)
\end{bmatrix}
\]

For this to work, of course, we need the matrix $H_s$ to be invertible (non-singular). This may not always be the case. We have seen in the previous demo how to compute matrix inverses and how to compute Hessians in Matlab if analytical expressions are used. In the case where the Hessian is computed numerically by finite differences we may use the approximation $2J^T_s J_s$.

**Problem 76.** The following model is to be used for data fitting

\[ y(x) = Ax^2 + Bx + 2Bx^{-1} \]

where $A$ and $B$ are parameters to be determined. In terms of the above model: (a) Derive the Jacobian matrix.

(b) Derive the design matrix.

(c) Derive the Hessian matrix.

(d) Derive the covariance matrix.

(e) Derive the normal equations for this model.
Chapter 9

Chi-square test

In this section we will describe the chi-square test. It can be used in at least two ways:

- In Regression Analysis: to determine the distance between the data and the fit. Here, the sum is over all data points.
- Test of Expected Distribution: here the test works on categorical data and we use it to determine the distance between two “histograms”. Here, the sum is over all bins of the histogram.

\[ \chi^2 \text{ test} \]

Suppose that we have \( \nu \) independent random variables \( X_i \) \((i = 1, \ldots, \nu)\) each normally distributed with mean \( \mu_i \) and variance \( \sigma_i^2 \). The chi-square is defined by:

\[
\chi^2 = \frac{(X_1 - \mu_1)^2}{\sigma_1^2} + \frac{(X_2 - \mu_2)^2}{\sigma_2^2} + \cdots + \frac{(X_\nu - \mu_\nu)^2}{\sigma_\nu^2} = \sum_{i=1}^{\nu} \frac{(X_i - \mu_i)^2}{\sigma_i^2} = \sum_{i=1}^{\nu} \tilde{R}_i^2
\]

Because the data is normally distributed, the normalized residuals are close to 1 (66% of all \( \tilde{R}_i \) values are within \( \pm 1 \) of 0). Hence, given a set of measurements \( \{X_i\} \) \((i = 1, \ldots, \nu)\), if we have chosen the \( \mu_i \) and \( \sigma_i^2 \) correctly, we may expect that a calculation of \( \chi^2 \) will be approximately equal to \( \nu \). If it is, then we may conclude that the data are well described by the values we have chosen for \( \mu_i \). If a calculated value of \( \chi^2 \) turns out to be much larger than \( \nu \), and we have correctly estimated the values for the \( \sigma_i^2 \), we may conclude that our data are not well described by our set of values \( \mu_i \).
9.2. \( \chi^2 \) distribution

\( \chi^2 \) is a random variable. With degree of freedom \( \nu \) we denote it as \( \chi^2_\nu \). We have seen that the sampling distribution of \( \chi^2 \) is the chi-square distribution:

\[
P(\chi^2 \leq \chi^2_\nu < \chi^2_\nu + d(\chi^2)) = p_\nu(\chi^2)d(\chi^2) = \frac{e^{-\chi^2/2}(\chi^2)^{(\nu/2)-1}}{2^{\nu/2}\Gamma(\nu/2)}d(\chi^2).
\]

if the expected frequencies are at least equal to 5. Here \( \Gamma \) is the gamma function:

\[
\Gamma(t) = \int_0^\infty x^{t-1}e^{-x}dx.
\]

The approximation improves for larger values. Boldface \( \chi^2_\nu \) denotes the random variable and \( \chi^2 \) is its value in the above expression. This is the \( \chi^2 \) distribution with parameter \( \nu \) (degrees of freedom). Here, \( 0 \leq \chi^2 < \infty \). \( \Gamma(p + 1) \equiv \int_0^\infty x^p e^{-x}dx \), \( \Gamma(p + 1) = p\Gamma(p) \), \( \Gamma(1/2) = \sqrt{\pi} \).

The mean of \( \chi^2_\nu \) is \( \nu \) and its variance is \( 2\nu \). This can be seen as follows. For the mean:

\[
\bar{\chi}^2 = \sum_{i=1}^{\nu} \frac{(X_i - \mu_i)^2}{\sigma_i^2} = \sum_{i=1}^{\nu} \frac{1}{\sigma_i^2}(X_i - \mu_i)^2 = \sum_{i=1}^{\nu} \frac{1}{\sigma_i^2}\sigma_i^2 = \nu
\]

where we have used the definition of the variance \( (X_i - \mu_i)^2 = \sigma_i^2 \). For the variance, \( var(\chi^2) \), we have

\[
var \left( \sum_{i=1}^{\nu} \frac{(X_i - \mu_i)^2}{\sigma_i^2} \right) = \sum_{i=1}^{\nu} \frac{1}{\sigma_i^4}var(X_i - \mu_i)^2 = \sum_{i=1}^{\nu} \frac{1}{\sigma_i^4} \left[ (X_i - \mu_i)^2 - \sigma_i^4 \right] = 2\nu
\]

where we have used the 4th central moment of a Gaussian distribution, \( (X_i - \mu_i)^4 = 3\sigma_i^4 \).

The \( \chi^2 \) distribution \( p_\nu(\chi^2) \) is plotted in Fig. 9.1, for different values of \( \nu \).

9.2.1. How to use the test. Suppose that we have \( k \) experimentally measured quantities \( X_i \) and we want to test whether they are well-described by some set of hypothesized values \( \mu_i \). We form a sum using Eq. 9.1. It will contain \( k \) terms, constituting a sample value for \( \chi^2 \).

The expected value of \( \chi^2 \) will not be \( n \) because the \( \{X_i\} \) may have been used to make estimates for the parameters of the model (e.g., \( \mu_i \) and \( \sigma_i^2 \)). Such relations (e.g., sample mean, sample variance) are said to reduce the number of degrees of freedom. If there are \( r \) such constraints then we number of degrees of freedom becomes:

\[
\nu = k - r
\]
and the resulting \( \chi^2 \) will be one having \( \nu \) (rather than \( k \)) degrees of freedom. Thus, the sum (Eq. 9.1) should be close to \( \nu \). In practice \( r \) will always be at least 1, because the data is normally used to estimate at least some parameter, such as the mean, variance, etc.

Generally speaking, \( \chi^2/\nu \approx 1 \) means the model represented by the \( \mu_i \) and \( \sigma^2_i \) is probably fine. If \( \chi^2/\nu \ll 1 \), we may conclude that either (i) the model is valid, but by chance, \( \chi^2 \) ended up being too small. (ii) we have overestimated the values of \( \sigma^2_i \). (iii) the data is fraudulent (too good to be true). If \( \chi^2/\nu \gg 1 \), we can conclude either that (i) the model is valid but by chance, \( \chi^2 \) ended up being too high. (ii) the model is so poorly chosen that an unacceptably large value of \( \chi^2 \) has resulted. (iii) data is not normally distributed. (Remember our assumption for least squares fitting that the data must be normally distributed about the model.)

9.2.2. Example 1: measuring mass of object using different types of balances. Suppose that we measure the mass of an object using four different balances (Cole-Parmer, Fischer, VWR and Nimbus). The results of the mass measurements are as follows:

<table>
<thead>
<tr>
<th>Detector (scale)</th>
<th>mass (g)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Cole-Parmer</td>
<td>91.161 ± 0.013</td>
</tr>
<tr>
<td>Fischer</td>
<td>91.174 ± 0.011</td>
</tr>
<tr>
<td>VWR</td>
<td>91.186 ± 0.013</td>
</tr>
<tr>
<td>Nimbus</td>
<td>91.188 ± 0.013</td>
</tr>
</tbody>
</table>

The listed uncertainties are estimates of the \( \sigma_i \), the standard deviations for each of the measurements. As can be seen, the error bars overlap. The question we would like to answer is: Can these data be well described by
a single number, namely an estimate of the mass made by determining the weighted mean of the four measurements?

The weighted mean and its error are:

\[
\bar{m} = \frac{\sum m_i / \sigma_i^2}{\sum 1 / \sigma_i^2} \quad \text{and} \quad \text{var}(\bar{m}) = \frac{1}{\left(\sum 1 / \sigma_i^2\right)^2} \sum \frac{1}{\sigma_i^2} \text{var}(m_i) = \frac{1}{\sum 1 / \sigma_i^2}.
\]

The value (for the 'true mean') we would report is:

\[
\bar{m} \pm \sqrt{\text{var}(\bar{m})} = 91.177 \pm 0.006.
\]

We then form \(\chi^2\):

\[
\chi^2 = \frac{\sum_{i=1}^{4} (m_i - \bar{m})^2}{\sigma_i^2} \approx 2.78.
\]

Here we have \(\nu = 4 - 1 = 3\) degrees of freedom (not 4), since we have used the weighted mean of the four measurements to estimate the value of the true mass, and this uses up one degree of freedom. The value \(\chi^2 / \nu = 2.78 / 3 \approx 0.93\) is close to 1, and so we conclude that the model is probably fine. In other words, the weighted mean (and assumption of Gaussian distribution of the measurements) is valid. Note: our claim that 0.93 is close to 1 should be decided based on a statistical significance criterion (level of confidence). We will see later how to do this.

### 9.3. Test of Expected Distribution

The test of expected distribution (also called the \(\chi^2\) test) is used to test the nature of a statistical distribution from which some random sample is drawn. Suppose that we have \(k\) classes (e.g. designated categories or bins in a histogram), with probabilities \(p_1, p_2, \ldots, p_k\) assigned to each class. If the chosen classes account for all the data, then \(\sum p_i = 1\). If we take the data and plot a histogram, we classify the data: this is done by counting the number of observation falling into each of the \(k\) classes (or bins). We have \(n_1\) counts in the first class, \(n_2\) counts in the second class, and so on, up to \(n_k\) counts in the \(k\)th class. Suppose there is a total of \(N\) observations, so \(\sum n_i = N\).

Then it can be shown that the sum

\[
\sum_{i=1}^{k} \frac{(n_i - Np_i)^2}{Np_i}
\]

has approximately the \(\chi^2\) distribution with \(\nu = k - r\)
9.3. Test of Expected Distribution

degrees of freedom, where \( r \) is the number of constraints, or relations used to estimate the \( p_i \) from the data. \( r \) will always be at least 1, since it must be that \( \sum n_i = \sum Np_i = N \sum p_i = N \). In other words, \( \nu \) is given by:

- \( \nu = k - 1 \) if the expected frequencies can be computed without having to estimate the population parameters from sample statistics. We subtract 1 from \( k \) because of the constraint condition \( \sum_j E_j = N \), which states that if we know \( k - 1 \) of the expected frequencies, the remaining frequency can be determined.
- \( \nu = k - 1 - m \) (i.e. \( r = m + 1 \)) if the expected frequencies can be computed only by estimating \( m \) population parameters from sample statistics.

Since \( Np_i \) is the mean, or expected value of \( n_i \), the form of \( \chi^2 \) corresponds to summing, over all classes, the squares of the deviations of the observed \( n_i \) from their mean values divided by their mean values. The above formula can be rewritten, using \( E_i = Np_i \) and \( O_i = n_i \):

\[
\chi^2 = \sum_{i=1}^{k} \frac{(O_i - E_i)^2}{E_i}
\]

where \( O_i \) stands for “observed value” and \( E_i \) stands for “expected value”.

9.3.1. The case of Poisson counts. We note that Eq. (9.3) looks different from (9.1). However, in the case of Poisson distributed values, the mean is equal to the variance. So this link with the Poisson distribution provides the motivation for the definition (9.3). In fact, Eq. (9.4) is obtained from Eq. (9.1) by taking \( \sigma_i^2 = E_i \) (the variance of the Poisson distribution is also the mean). This is justified because the vertical axis in a histogram is the number of counts whereas the number of counts is approximately Poisson distributed.

Equation (9.3) is sometimes written as the sum of observed minus expected values squared (divided by expected value). Your textbook uses \( O_i \) to denote the number of counts observed (during the \( i \)th interval or as a frequency of occurrence for some category/bin). If \( E_i \) denotes the expected number of counts (i.e. the mean value of \( O_i \)), we write:

\[
\chi^2 = \sum_{i=1}^{k} \frac{(O_i - E_i)^2}{E_i} \quad \text{(Poisson statistics)}
\]

\(^1\)We have seen, in the case of Poisson statistics, that the error is given by \( \sqrt{O_i} \). This approach is problematic for low counts, as low counts lead to large random fluctuations of the error in the denominator of the chi-square. One solution is to re-bin the data such that the counts are not too low.
$O_i$: observed count/frequency,
$E_i$: expected count (model), as derived from the Poisson distribution formula. Note: $E_i$ is not necessarily equal to $P(n; \pi) = e^{-\pi} \pi^n / n!$; it depends on how the data is binned. However, it will be derived using $P(n; \pi)$.

$k$: number of bins/categories

For good fits, the error obeys $\sqrt{E_i} \approx \sqrt{O_i}$. If

$$\chi^2 \gg \nu = k - r$$

($r$: number of parameters in the distribution or the number of constraints used to compute the expected values) there is significant disagreement between observed and expected distribution. Please note that because $\chi^2$ itself must be dimensionless, this test only makes sense if $O$ and $E$ are dimensionless quantities, such as counts.

Because $E_i$ and $O_i$ are counts (frequencies of occurrence) rather than physical quantities (with units), you must bin your data into a histogram and then compare the histogram to the distribution to which it is fitted to. In practice, the distribution needs to use experimental data to compute its parameters (e.g. obtained via maximum likelihood estimation). When such parameters are computed from the data, we say that the degrees of freedom have been reduced. This is why the number of degrees of freedom is equal to the number of categories minus the number of parameters (or more precisely, the number of conditions used to estimate the parameters of the distribution from the data), $\nu = k - r$.

9.3.2. Example 2: counting trees on acres of land. Suppose that you own a large parcel of land in the desert that is divided into 48 acres (a rectangular region of 6 acres by 8 acres), and you hire someone to count the trees in each acre. The data ranges from 0 trees up to 6 trees per acre. But because there are very few acres that contain 4 or more trees, you decide to lump them all together into the same category (“4 or more trees”). You find the following data:

<table>
<thead>
<tr>
<th>Observed counts</th>
<th>1 2 3 4 5</th>
</tr>
</thead>
<tbody>
<tr>
<td>Category</td>
<td></td>
</tr>
<tr>
<td>Number of trees in an acre</td>
<td>0 1 2 3 ≥4</td>
</tr>
<tr>
<td>Frequency (number of acres in this category)</td>
<td>9 9 10 14 6</td>
</tr>
</tbody>
</table>

Suppose that for bin 5, the six counts were actually 4, 4, 5, 5, 6, 6. You want to check whether or not the counting statistics can be described by a Poisson distribution. The Poisson distribution requires the estimation of 1

\footnote{Or number of data points, $n$, if the test is applied as in the case of regression analysis, as discussed in the previous section).}
parameter. The maximum likelihood estimator of the Poisson parameter is the arithmetic average (mean) of the sample data:

\[
\bar{n} = \frac{(0)(9) + (1)(9) + (2)(10) + (3)(14) + (4)(2) + (5)(2) + (6)(2)}{9 + 9 + 10 + 14 + 2 + 2 + 2} = 2.10.
\]

We have 48 counts and the expected number of counts in the \(j\)th category (bin) is given by \(48 \times p_j\), where \(p_j\) is the probability of observing the stated counts in this category. The probabilities \(p_j\) are obtained using the Poisson distribution. Let \(X\) be the random variable that counts the number of trees in an acre. The probabilities are:

\[
p_1 = P(X = 0) = e^{-2.10} = e^{-2.10},
\]

\[
p_2 = P(X = 1) = \frac{e^{-2.10}(2.10)^1}{1!} = e^{-2.10} \cdot 2.10,
\]

\[
p_3 = P(X = 2) = \frac{e^{-2.10}(2.10)^2}{2!} = e^{-2.10} \cdot 2.205,
\]

\[
p_4 = P(X = 3) = \frac{e^{-2.10}(2.10)^3}{3!} = e^{-2.10} \cdot 1.5435,
\]

\[
p_5 = P(X \geq 4) = 1 - P(X \leq 3) = 1 - e^{-2.10}(1 + 2.10 + 2.205 + 1.5435) = 7.74.
\]

Then, the expected frequencies/counts in each category/bin are:

\[
E_1 = 48 \cdot p_1 = 5.8779,
\]

\[
E_2 = 48 \cdot p_2 = 12.3436,
\]

\[
E_3 = 48 \cdot p_3 = 12.9608,
\]

\[
E_4 = 48 \cdot p_4 = 9.0726,
\]

\[
E_5 = 48 \cdot p_5 = 7.74.
\]

The test of expected distribution (\(\chi^2\)) can then be calculated:

\[
\frac{(9 - 5.8779)^2}{5.8779} + \frac{(9 - 12.3436)^2}{12.3436} + \frac{(10 - 12.9608)^2}{12.9608} + \frac{(14 - 9.0726)^2}{9.0726} + \frac{(6 - 7.74)^2}{7.74},
\]

which equals to 6.3097. This number should be compared to the degrees of freedom, \(\nu = k - r = 5 - 1 = 4\). Is this a good fit? If the answer is yes, then we say that the data is well described by a Poisson distribution with parameter \(\bar{n} = 2.10\). To determine this, we need to check whether \(\chi^2/\nu \approx 1\) according to some level of confidence (see Section 9.3.4).

9.3.3. Example 3: height of people. Let us use the chi-square test to test whether a data sample consisting of the heights of \(N=66\) people can be assumed to be drawn from a Gaussian distribution or not. We first arrange the data in the form of a frequency distribution, listing for each height \(h\), the value of \(n(h)\), the number of people in the sample whose height is \(h\) (\(h\) is in inches):
We make the hypothesis that the heights are distributed according to the Gaussian distribution, namely that the probability \( p(h) \, dh \) that a height falls between \( h \) and \( h + dh \) is given by:

\[
p(h) \, dh = \frac{1}{\sqrt{2\pi\sigma^2}} e^{-(h-\mu)^2/2\sigma^2} \, dh.
\]

This expression, if multiplied by \( N \), will give, for a sample of \( N \) people, the theoretically expected number of people \( n_e(h) \, dh \) whose height should be between \( h \) and \( h + dh \):

\[
n_e(h) \, dh = \frac{N}{\sqrt{2\pi\sigma^2}} e^{-(h-\mu)^2/2\sigma^2} \, dh.
\]

In our example, \( N = 66 \). In our table we have grouped the data into bins, each of size 1 inch. A useful approximation is to take \( dh = 1 \) inch, which gives the expected number of people having a height \( h_j \):

\[
(9.5) \quad n_e(j) = \frac{N}{\sqrt{2\pi\sigma^2}} e^{-(h_j-\mu)^2/2\sigma^2}.
\]

The sample mean \( \bar{h} \) and the sample standard deviation \( s \) are our best estimates of \( \mu \) and \( \sigma \). We find, calculating from the data:

\[
\bar{h} = 64.9 \text{ inches} \quad \text{and} \quad s = 2.7 \text{ inches}.
\]

Using these values we may calculate from Equation (9.5), the number expected in each bin, which the following results:

\[
\begin{array}{cccccccccccccc}
  h & 58 & 59 & 60 & 61 & 62 & 63 & 64 & 65 & 66 & 67 & 68 & 69 & 70 & 71 & 72 & 73 \\
  n(h) & 1 & 0 & 1 & 4 & 6 & 7 & 13 & 8 & 11 & 2 & 7 & 4 & 1 & 0 & 0 & 1 \\
  n_e(j) & 0.3 & 0.9 & 1.8 & 3.4 & 5.5 & 7.6 & 9.3 & 9.9 & 9.1 & 7.3 & 5.0 & 3.1 & 1.6 & 0.7 & 0.3 & 0.1
\end{array}
\]

In applying the chi-square test to a situation of this type, it is advisable to re-group the data into new bins (classes), such that the expected number occurring in each bin is greater than 4 or 5; otherwise the theoretical distributions within each bin become too highly skewed for meaningful results. Thus, in this situation we shall put all the heights of 61 inches or less into a single bin, and all the heights of 69 inches or more into a single bin. This groups the data into a total of 9 bins (or classes), with actual numbers and expected numbers in each bin given as follows (note the bin sizes need not be equal):

\[
\begin{array}{cccccccc}
  h & \leq 61 & 62 & 63 & 64 & 65 & 66 & 67 & \geq 69 \\
  n(h) & 6 & 6 & 7 & 13 & 8 & 11 & 2 & 7 & 6 \\
  n_e(j) & 6.5 & 5.5 & 7.6 & 9.3 & 9.9 & 9.1 & 7.3 & 5.0 & 5.8
\end{array}
\]
9.3. Test of Expected Distribution

Now we calculate the value of $\chi^2$ using these data:

$$\chi^2 = (6 - 6.5)^2 \frac{6.5}{5.5} + (6 - 5.5)^2 + \cdots + (6 - 5.8)^2 \frac{5.8}{5.5} = 6.96$$

Since we grouped our data into 9 classes, and since we have used up three degrees of freedom by demanding that: (i) the sum of the $n_j$ be equal to $N$, (ii) the mean of the distribution be equal to the sample mean and (iii) the variance is equal to the sample variance, the are 6 degrees of freedom left. Hence $\chi^2/\nu = 6.96/6 \approx 1.16$, which is close to 1. Therefore, we have no good reason to reject our hypothesis that our data are drawn from a Gaussian distribution function.

9.3.4. $\chi^2$ distribution and significance tests. The $\chi^2$ distribution is used to perform “significance tests”, as explained below by way of examples. The main idea is that, as we repeat our experiment and collect values of $\chi^2$, if our model is a valid one, these data will be clustered about the median value of $\chi^2_\nu$, with about half of them greater than the median and half less than the median. This median value, which we denote $\chi^2_{\nu,0.5}$ is determined by:

$$\int_{\chi^2_{\nu,0.5}}^{\infty} p_\nu(\chi^2) d\chi^2 = 0.5.$$

In other words, we expect that a single measured value of $\chi^2$ will have a probability of 0.5 of being greater than $\chi^2_{\nu,0.5}$. Generalizing this idea, to say that we expect a single measured value of $\chi^2$ will have a probability $\alpha$ of being greater than $\chi^2_{\nu,\alpha}$, where $\chi^2_{\nu,\alpha}$ is defined by

$$\int_{\chi^2_{\nu,\alpha}}^{\infty} p_\nu(\chi^2) d\chi^2 = \alpha.$$

This definition is illustrated by the inset of Fig. 9.2, which plots $\alpha$ versus the normalized chi-square, $\chi^2_\nu/\nu$. $\alpha$ is the probability that a sample chi-square will be larger than $\chi^2_\nu$, as shown in the inset. Each curve is labeled by $\nu$, the number of degrees of freedom.

To use this test, we calculate $\chi^2$ and $\nu$. This gives the value $\chi^2/\nu$. Choose $\alpha$, the significance level. Use the above figure to determine the corresponding value of $\chi^2_{\nu,\alpha}/\nu$. Compare this value with our sample value $\chi^2/\nu$. If we find that $\chi^2/\nu > \chi^2_{\nu,\alpha}/\nu$, then either: (i) a statistically improbable excursion of $\chi^2$ has occurred; (ii) our model is poorly chosen or (iii) the data is not normally distributed. If $\chi^2/\nu < \chi^2_{\nu,\alpha}/\nu$, we have either: (i) a valid model, but statistically improbable excursion of $\chi^2$, (ii) the data is fraudulent (“too good to be true”). Note: a poor model can only increase the value of $\chi^2$. 
• **Example 1: measuring mass of object.** We expect this value of \( \chi^2 \) to be drawn from a chi-square distribution with 3 degrees of freedom. From the above graph of \( \alpha \) vs \( \chi^2 \nu \) for \( \nu = 3 \), the value \( \chi^2 / \nu = 2.78 / 3 \approx 0.93 \) corresponds to \( \alpha \) of about 0.42. This means that if we were to repeat the experiments we would have about a 42 percent chance of finding a \( \chi^2 \) for the new measurement set larger than 2.78, assuming that the hypothesis (of a uniform distribution) is correct. We have therefore no good reasons to reject the hypothesis and conclude that the four measurements of the mass are consistent with each other. We would have had to have found \( \chi^2 \) in the vicinity of 8 (leading to an \( \alpha \) of about 0.05) to have been justified in suspecting the consistency of the measurement.\(^3\) The fact that our value of \( \chi^2 / 3 \) is close to 1 is reassuring.

• **Example 2: counting trees on acres of land.** Let us return to the example of counting trees. There were 5 bins less 2 degrees of freedom (\( \nu = 5 - 2 = 3 \)): (1) one degree of freedom was used to constrain the total number of acres; (2) the second degree of freedom was used to estimate the parameter of the Poisson distribution. Suppose that the confidence level is 5\%, i.e. we set \( \alpha = 0.05 \). The critical value of the test is \( \chi^2_{3,0.05} = 7.8 \). (The above graph, at \( \alpha = 0.05 \) and \( \nu = 3 \) gives a value of \( \chi^2 / \nu \sim 2.6 \).) Comparing the value of \( \chi^2 = 6.3 \) we calculated in the example, this value is less than the critical value of the test. Thus, at the 0.05 level of significance

\(^3\)From the graph, at \( \alpha = 0.05 \) and \( \nu = 3 \) the value of \( \chi^2_{\alpha, \nu} = \chi^2_{3,0.05} / 3 \) is 2.6 (i.e., \( \chi^2 \sim 7.8 \)). The fact that 0.9 is less than 2.6 does not mean that we have a poor model. (For that we would need a value greater than 2.6.) It could be, for example, that the data describe a statistically improbable excursion of \( \chi^2 \).
we fail to reject the hypothesis that the data is well represented by the Poisson distribution with the estimated parameter, \( \bar{n} = 2.10 \).

- **Example 3: height of people.** Here \( \chi^2/\nu = 6.96/6 \approx 1.16 \), leading to an \( \alpha \) of about 0.33. Therefore, we have no good reason to reject our hypothesis that our data are drawn from a Gaussian distribution function. (From the graph, \( \chi^2_{\nu,0.05}/\nu \) for \( \nu = 6 \) is approximately equal to 2.2; and since 1.16 is less than 2.2, we cannot reject the model.)

**9.4. Problem**

**Problem 77.** The middle A note on a piano is normally tuned to the nominal value of 440 Hz. The piano you bought was delivered to your home and tuned properly (440 Hz). You monitor over time the exact pitch of the middle A note (normally tuned to 440 Hz). Of course, the piano progressively detunes over time and the pitch changes. The instrument you use to measure the pitch (some app on your phone) has its own error bar, which may fluctuate over time (for some strange reason). Think of this pitch as some kind of random variable \( X \), where \( X = 440 \) at time zero, after the piano was properly tuned, but taking other values at other times.

You measure the pitch at day 1 (1 day after tuning), day 40, day 80 and day 120. The results are:

<table>
<thead>
<tr>
<th>Day you measure actual pitch</th>
<th>pitch of A note (Hz)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Day 1</td>
<td>440.61 ± 15.11</td>
</tr>
<tr>
<td>Day 40</td>
<td>433.74 ± 0.15</td>
</tr>
<tr>
<td>Day 80</td>
<td>432.86 ± 0.10</td>
</tr>
<tr>
<td>Day 120</td>
<td>432.88 ± 0.13</td>
</tr>
</tbody>
</table>

(a) What is the mean of \( X \), as calculated by the data and its error bars?

(b) Calculate the uncertainty of the mean calculated in (a).

(c) Report the pitch \( A \) as best value \( \pm \) uncertainty, with the correct number of significant figures.

(d) Let’s assume that the measurements are Gaussian distributed about their mean value. Is this a good assumption? (i.e. “prove” or disprove it, using the appropriate statistical test)

(e) Why would anyone assume that a measurement should be Gaussian-distributed?
Solution. (a) Use weighted average:

\[ X = \frac{\sum x_i / \sigma_i^2}{\sum 1 / \sigma_i^2} = 433.0581 \text{ Hz} \]

(b)

\[ \text{var}X = \frac{1}{\sum 1 / \sigma_i^2} = 0.049 \]

The uncertainty is the square root of this (0.0701 Hz).

(c) Keeping 1 digit for the error: 433.06 ± 0.07 Hz

(d) We use the chi square test:

\[ \chi^2 = \sum_{i=1}^{4} \frac{(x_i - X)^2}{\sigma_i^2} = 26.72 \]

the number of degrees of freedom is 4-1=3. Dividing by 3 gives 8.9, which is much larger than 1. So Gaussian may not be a good assumption.

(e) Per central limit theorem (CLT). However, CLT does not always apply.

Problem 78. You arrive (by car) at the Dodgers stadium for a ball game and want to park near the entrance. However, the parking attendant redirects traffic into two threads. The first thread leads cars away from the stadium. The second thread leads to parking spots near the stadium’s entrance. You were redirected far away from the stadium. You suspect that the parking attendant is being biased against ugly cars (such as yours) and set out to test the hypothesis that the attendant’s policy is to redirect more luxury cars toward the stadium and keep the ugly looking cars away from the stadium.

(a) In order to test the hypothesis that the parking attendant was fair, you ask your friend to stand nearby and record the number of times that luxury cars get redirected toward the stadium. Suppose that out of 500 observations, he observed 265 luxury cars and 235 ugly cars were redirected toward the stadium. Is the attendant fair or biased?

(b) After parking your car far out and walking nearly 1/2 mile back to the stadium, on your way back, you walk by the attendant and decide whether he is biased or not. Before making this decision, you want to double check the results from part (a). So you ask your friend to go and repeat this experiment. Out of 500 observations, your friend obtains an even higher count of luxury cars (270 luxury cars and 230 ugly cars). Is the attendant biased?
Solution. (a) 
\[ \frac{(265 - 250)^2}{250} + \frac{(235 - 250)^2}{250} = 1.8 \]
because this number is close to 1, we cannot conclude there is bias. (Degrees of freedom: 2-1, since there are two categories, and 1 constraint for the total number of cars.)

(b) 
\[ \frac{(270 - 250)^2}{250} + \frac{(230 - 250)^2}{250} = 3.2 \]
again, this number is not orders of magnitude different than 1, so we cannot conclude there is bias. ■
10.1. Basics of the MATLAB command environment

MATLAB is a scientific computing software platform that boasts millions of users worldwide. Its broad user base includes chemists, physicists, computer scientists, engineers (including finance), life scientists and mathematicians. It allows you to run tasks by executing scripts (lists of commands) or to do computations one at a time by typing commands in its console. In the MATLAB environment you can store variables (scalars, vectors, matrices, strings, structures, etc.) whereas MATLAB itself features thousands of different commands that you can use. Variables can be created by setting their value. For example:

\[
a = 1
\]

\[
a = 1
\]
This command creates a new variable \( a \) (if it does not exist already), and sets its value equal to 1. To check that you have created this variable, type:

\[
\text{whos}
\]

Arrays of numbers can be stored as matrices or vectors. The following command creates a row vector:

\[
a=[4\ 7\ 10\ 12\ 18\ 8\ 7\ 6\ 4.5\ 2.7\ 13.8]
\]

\[
a =
\]

Columns 1 through 9

\[
\begin{array}{cccccccc}
4.0000 & 7.0000 & 10.0000 & 12.0000 & 18.0000 & 8.0000 \\
7.0000 & 6.0000 & 4.5000 \\
\end{array}
\]

Columns 10 through 11

\[
\begin{array}{cccc}
2.7000 & 13.8000 \\
\end{array}
\]

This could be, for example, a set of experimental values. MATLAB has many internal commands that are available to us for purposes of doing computations. For example, the standard deviation of the entries stored in \( a \) can be computed using the \texttt{std} command:

\[
\text{std(a)}
\]

\[
\text{ans} =
\]

\[
4.6228
\]

Internal MATLAB commands are scripts (lists of commands) that can be viewed. For example, entering:

\[
\text{type std}
\]

we get:

\[
\begin{verbatim}
function y = std(varargin)
%STD Standard deviation.
% For vectors, \( Y = \text{STD}(X) \) returns the standard deviation. For ...
matrices,
\end{verbatim}
\]
As you can see, there is only 1 command that does anything: \( y = \sqrt{\text{var(varargin{:})}} \). It calls another command, \texttt{var}, to calculate the sample variance. Then it takes the square root of this variance and returns it as the standard deviation. You can view the code for \texttt{var} by typing \texttt{type var}.
The mean is calculated as follows:

```matlab
mean(a)
```

```matlab
ans =
8.4545
```

Likewise, there are commands for calculating the median:

```matlab
median(a)
```

```matlab
ans =
7
```

There is also a command to calculate percentiles. Type `help prctile` for more information. For example:

```matlab
prctile(a,50)
```

```matlab
ans =
7
```

which is the median. Also,

```matlab
prctile(a,10)
```

```matlab
ans =
3.4800
```

Vector addition:

```matlab
a=[1 2 3]
a =
1 2 3
```

```matlab
b=[4 5 6]`
10.1. Basics of the MATLAB command environment

\[
b = \\
\begin{bmatrix}
4 & 5 & 6 \\
\end{bmatrix}
\]

\[
a+b
\]

\[
\text{ans} = \\
\begin{bmatrix}
5 & 7 & 9 \\
\end{bmatrix}
\]

Multiplication of vector (\(a\)) by a scalar (2):

\[
a
\]

\[
a = \\
\begin{bmatrix}
1 & 2 & 3 \\
\end{bmatrix}
\]

\[
2*a
\]

\[
\text{ans} = \\
\begin{bmatrix}
2 & 4 & 6 \\
\end{bmatrix}
\]

Matrices can be created by separating the different rows using semicolons:

\[
>> a=[1 \ 2 \ ; \ 3 \ 4 ]
\]

\[
a = \\
\begin{bmatrix}
1 & 2 \\
3 & 4 \\
\end{bmatrix}
\]

\[
b=[5 \ 6 \ ; \ 7 \ 8]
\]

\[
b = \\
\begin{bmatrix}
5 & 6 \\
7 & 8 \\
\end{bmatrix}
\]

\[
\text{whos}
\]
Matrix addition (element by element addition):

\[ \begin{bmatrix} 2 & 4 \\ 8 & 16 \end{bmatrix} + \begin{bmatrix} 3 & 6 \\ 9 & 18 \end{bmatrix} = \begin{bmatrix} 5 & 10 \\ 17 & 34 \end{bmatrix} \]

Matrix multiplication:

\[ \begin{bmatrix} 2 & 4 \\ 8 & 16 \end{bmatrix} \times \begin{bmatrix} 3 & 6 \\ 9 & 18 \end{bmatrix} = \begin{bmatrix} 30 & 72 \\ 144 & 336 \end{bmatrix} \]

For pointwise (element by element) multiplication, we use \( .* \) instead of \( * \):

\[ \begin{bmatrix} 2 & 4 \\ 8 & 16 \end{bmatrix} .* \begin{bmatrix} 3 & 6 \\ 9 & 18 \end{bmatrix} = \begin{bmatrix} 6 & 24 \\ 72 & 288 \end{bmatrix} \]

We can even take the elements of \( a \) and elevate them to the powers of the elements of \( b \):

\[ \begin{bmatrix} 2 & 4 \\ 8 & 16 \end{bmatrix} .^* \begin{bmatrix} 3 & 6 \\ 9 & 18 \end{bmatrix} = \begin{bmatrix} 8 & 64 \\ 4096 & 65536 \end{bmatrix} \]

Let's now multiply a 2 × 2 matrix by a 2 × 1 (column) vector:
10.1. Basics of the MATLAB command environment

c=[6 ; 8]
c =
   6
   8
The multiplication of a and c is:

a*c
ans =
   22
   50
We can take the natural log of the elements of a:

log(a)
ans =
   0
   0.6931
  1.0986
  1.3863
log means natural log. To get log base 10, there is a command log10. Or
use the property log_a b = log(b)/log(a).

Suppose you created a row vector,

c=[6 8]
c =
   6
   8
when it reality you needed a column vector. You can use the transpose
command to do that:

transpose(c)
ans =
   6
A shorthand is to add a prime after c:

```matlab
>> c'
ans =
   6
   8
```

If you try to multiply a and c you get an error message:

```matlab
>> a*c
Error using *  
Inner matrix dimensions must agree.
```

because c must be a column vector, not a row vector. Instead, what works is:

```matlab
>> a*transpose(c)
ans =
   22
   50
```

or

```matlab
>> a*c'
ans =
   22
   50
```

Ranges can be created by using the colon:

```matlab
>> 1:10
ans =
     1     2     3     4     5     6     7     8     9    10
```
If a different step size than 1 is desired, sandwich the step size between two colons:

```
1:0.5:10
```

ans =

```
Columns 1 through 9
1.0000  1.5000  2.0000  2.5000  3.0000  3.5000
   4.0000  4.5000  5.0000

Columns 10 through 18
  5.5000  6.0000  6.5000  7.0000  7.5000  8.0000
      8.5000  9.0000  9.5000

Column 19

10.0000
```

If you don’t know the step size but know how many points are needed, use `linspace`:

```
linspace(1,10,5)
```

ans =

```
1.0000  3.2500  5.5000  7.7500  10.0000
```

### 10.1.1. Taylor series approximation to the matrix exponential.

The exponential of a matrix $M$ is defined by its Taylor series expansion:

$$
\exp(M) = \sum_{j=0}^{\infty} \frac{M^j}{j!}.
$$

We can use MATLAB to see how this Taylor series converges to $\exp(M)$. Let’s compute partial sums (truncated Taylor series) and compare the results as the number of terms is increased. We will compare the results to the `EXPM` command in MATLAB, which uses the scaling and squaring algorithm (with Padé approximant) to compute the matrix exponential:

```matlab
>> help expm
```
expm Matrix exponential. expm(A) is the matrix exponential of A and is computed using a scaling and squaring algorithm with a Padé approximation.

Although it is not computed this way, if A has a full set of eigenvectors V with corresponding eigenvalues D then [V,D] = EIG(A) and expm(A) = V*diag(exp(diag(D)))/V.

EXP(A) computes the exponential of A element-by-element.

See also logm, sqrtm, funm.

The following code below will compute the partial sum up to \( j = NT \) (a value which you can set manually), for a 2 \( \times \) 2 matrix \( m \):

```
m=[1 2; 3 4]; % matrix to be exponentiated
NT=100;
w=eye(2);
for j=1:NT,
w = w + (m^j)/factorial(j);
end;
disp('Result:');
w disp('True value:');
expm(m)
disp('Distance:');
d=w-expm(m);
dis=sum(abs(d(:)))/4
```

In order to assess the convergence of the series, we repeat the calculation for several different values of \( NT \) (from 1 to 100) and measure the distance of the results (\( w \)) from the true value (\( \text{expm}(m) \)), using the distance metric:

\[
d = \sum_{ij} |w_{ij} - \text{expm}(m)_{ij}|.
\]

where the sum runs over all matrix elements. Other metrics are possible. A plot of \( d \) versus \( NT \) is shown below:
We conclude that only about 10 terms are needed for convergence and calculation of the infinite series is not required. The exact number of terms needed depends on the desired precision.

10.1.2. Plotting histograms: Illustration of the Central Limit Theorem. The command in MATLAB to plot histograms is `hist` on earlier versions of MATLAB and `histogram` in the later versions. Another useful function is `histfit`, which will not only plot a histogram, but also will fit a Gaussian PDF to it. Let’s look at the central limit theorem (CLT).

We know the CLT predicts that the sum of random variables (regardless of their distribution, as long as they have finite mean and variance) eventually converges to a normal distribution in the limit of large $n$:

$$\frac{X_1 + \cdots + X_n}{n} \rightarrow N(\mu, \sigma^2 / n) \quad \text{as } n \rightarrow \infty.$$ 

In the code below, $NA$ uniformly distributed random variables are added together. The results for $NA=1,2,3,100$ are shown in Fig 10.1. As can be seen, the $n = 1$ case is uniform, the $n = 2$ case yields a tent function, whereas $n = 3$ looks increasingly more like a Gaussian. The $n = 100$ case looks Gaussian. Notice also the width of the distribution narrows. This is due to the variance of the arithmetic average, $\sigma^2 / n$, which decreases as $1 / n$.

```matlab
v=zeros([1 10000]);
NA=100; % of rv's to be added in sum
for j=1:NA,
    v=v+rand([1 10000]); % randomly generate 10,000 pts
end; % and add them together (sum NA such rv's)
v=v/NA; % normalize to the # of rv's added
figure; histfit(v,20); % plot histogram, fit a gaussian to it
set(gca,'xlim',[0 1]);
set(gca,'fontsize',16);
```

When dealing with summations of \( n \) terms, it is often useful to look at the partial results of the sum as a way to study convergence of the sum. To this end, the command `cumsum` can be used. The code below illustrates the example of computing the following integral numerically, using a random number generator. Let’s compute the average of \( \sin(x) \) from 0 to \( 2\pi \):

\[
I = \frac{1}{2\pi} \int_0^{2\pi} \sin(x)dx.
\]

This integral is of the form \( \int f(x)p(x)dx \), where \( p(x) = (2\pi)^{-1}1_{[0,2\pi]}(x) \) is the PDF for the uniform distribution over the interval \([0, 2\pi]\). Thus, it is the expectation value \( \mathbb{E}_p(\sin(X)) \), which can be approximated by the law of large numbers (LLN) as a sum:

\[
I = \mathbb{E}_p(\sin(X)) \approx \frac{1}{n} \sum_{i=1}^{n} \sin(X_i), \quad X_i \sim p(x).
\]

Thus, \( I \) can be approximated by generating uniformly distributed random numbers \( \{X_i\} \) and taking the arithmetic average of the \( \{\sin(X_i)\} \). The code below generates the summation for different \( n \).
10.1. Basics of the MATLAB command environment

**Figure 10.2.** Illustration of the law of large numbers. Plot of \( \frac{1}{n} \sum_{i=1}^{n} \sin(X_i) \) vs \( n \), from \( n = 1, \ldots, 10000 \) (horizontal axis). As \( n \) increases, the value of the sum approaches 0, the value of the integral.

```matlab
NP=10000; % max # of points in LLN sum (n=1,...,NP)
x=(2*pi)*rand([1 NP]); % sample integrand uniformly from 0 to 2*pi
sx=(1/2/pi)*sin(x); % function to be integrated
NL=linspace(1,NP,NP); % normalize to # of pts. in sum
ysx=cumsum(sx)./NL;
figure; plot(ysx); % convergence of sum vs n
set(gca,'fontsize',16);
```

10.1.4. Fitting a trendline to data plotted in a figure window. Data can be analyzed using the basic curve fitting tools available from the pull-down menus of a figure window. To show how this works, let us start by creating a fake data set. The horizontal axis is:

```matlab
t=1:100;
```

Create a slope:

```matlab
slope=10;
```

This command will create a straight line graph, with gaussian noise added:

```matlab
f=t*slope + 60*randn([1 100]);
```

To get more information about random numbers type `help randn`. Plot the results as follows:

```matlab
figure; plot(t,f,'o');
```
In the next section we shall analyze this data, as well as data from the following function (straight line plus quadratic component):

```matlab
g = f + 0.1*t.^2; figure; plot(t,g,'o');
```

and this one (straight line plus sinusoidal component):

```matlab
h = f + 50*sin(0.2*t); figure; plot(t,h,'o');
```

The straight line data looks like this:

Basic curve fitting can be done by choosing Tools in the Figure menu, followed by Basic Fitting. If we expect the data to be described by a straight line, select linear from the menu. MATLAB will plot the trendline and list the results from the fit in a box.
If you check the box Plot residuals, MATLAB will divide the window into two sections and add a plot of the differences between the trendline and the fitted curve. Residuals are a good way to see if the model chosen (in this case, a straight line) is suitable for describing the experimental data. The residuals are shown in the above example. If the data is a good fit to the model, the residuals should be normally distributed (Gaussian) with mean 0.

More formally, the $i$-th residual $r_i$ is defined as:

$$ r_i = y(x_i) - y_i $$

where $y(x_i)$ is the model and $y_i$ is the experimentally-measured data point. MATLAB also returns a quantity called Norm of residuals. MATLAB defines this quantity as

$$ \text{Norm of residuals} = \sqrt{\sum_{i=1}^{N} r_i^2} $$

where $N$ is the number of data points. This provides a measure of the goodness-of-fit. A perfect fit has zero residuals and the Norm of residuals will also be zero. A model for the fit should be chosen as to minimize the Norm of residuals. Choosing any model for the sake of minimizing Norm of residuals is generally not a good strategy.

The following link contains useful information on interactive data fitting


An example of poor data fitting is shown below. The best judge for this is visual inspection: the quadratic model clearly isn’t a good match to the data. The poor fit is also evidenced by the large value of Norm of residuals.
A bad choice of fitting function is shown below where a 7th order polynomial was used to fit the above linear data set. The fit to the 7th order polynomial gives a better value for the Norm of residuals. But it is also completely unphysical and yields no useful information about the experiment.

In general, you want to pick a model based on the physics of the system you are studying. You also should pick the simplest model possible, i.e. one with fewest parameters. As you increase the number of fitting parameters (i.e. by choosing more complicated functions), the fit look better (visually or according to Norm of residuals), but the model is meaningless from a physical standpoint.
To re-iterate this point: when fitting data, we want to minimize the residuals, but this should not be done at the expense of choosing an inadequate model. Start from the simplest model and build up gradually in complexity as needed. Examples are shown below.

In the following example, a noisy data set is fit well by a straight line. Inspection of the residuals show that the fit does not appear to be missing any underlying trends, i.e. the residuals are just noise and don’t exhibit any particular structure.

The second example below shows a similar data set, but one in which the linear fit leaves out residuals. Inspection of the structure of these residuals suggests that we are missing a quadratic component. You would then be better off to fit the data using a second order polynomial.

In such cases, you are left with the task of explaining the physical origin of this quadratic dependence. In the case of Hooke’s law, one possible explanation would be the over-stretching of the spring which is known to lead to non-linear behavior. Each experimental situation is different and must be examined in its own light.
The following interactive fit shows that a quadratic model leaves residuals that are just noise (no specific underlying structure). As an exercise, we could plot the histogram of these residuals and check that they are distributed normally with mean 0.

The final example is an experimental data set which is mostly a straight line, but where the residuals reveal an underlying sinusoidal behavior.
10.1.5. Can residuals plots help identify systematic errors? We remark that plotting the residuals can be a useful method for identifying systematic errors in your experiment. When the laws of physics state that the behavior should be modeled by a particular set of equations, but that the residuals of the fit to this model point to additional trends in the data, this is a sign that there may be systematic errors.

For example, a sinusoidal trend in the residuals that does not follow from the expected model may be suggestive of the presence of vibrations in your experimental apparatus. In this case, systematic errors could be mitigated by fitting the data to the original model, plus a sinusoidal term. The parameters of the model are extracted and those of the sinusoidal fit are discarded. A brute-force, but perhaps costly, way to reduce systematic errors is to improve the physical apparatus.

A careful analysis of the origins of systematic errors may enable you to identify and subtract these errors without having to re-design the experiment.

10.2. cftool – a GUI-based curve fitting tool

The cftool command in MATLAB is part of the Curve Fitting toolbox and provides a graphical user interface (GUI) for data fitting. In this section, we will demonstrate its use.

Let us first load a data set from the file noise.mat. The variable signoise contains a complex-valued vector which is the free induction decay (FID) from a nuclear magnetic resonance (NMR) experiment done on a test tube of water (a single resonance). The last two commands below will plot the signal in the time domain (the FID) and its Fourier transform (the NMR spectrum).
load noise
whos

<table>
<thead>
<tr>
<th>Name</th>
<th>Size</th>
<th>Bytes</th>
<th>Class</th>
<th>Attributes</th>
</tr>
</thead>
<tbody>
<tr>
<td>addnoise</td>
<td>207x1</td>
<td>3312</td>
<td>double</td>
<td>complex</td>
</tr>
<tr>
<td>multnoise</td>
<td>1x951</td>
<td>15216</td>
<td>double</td>
<td>complex</td>
</tr>
<tr>
<td>signoise</td>
<td>1x207</td>
<td>3312</td>
<td>double</td>
<td>complex</td>
</tr>
</tbody>
</table>

figure;plot(abs(signoise));
figure;plot(abs(fftshift(ifft(signoise))));

The result is:

![Free induction decay (FID) - time domain NMR signal](image1)

![NMR spectrum (Fourier transform of FID)](image2)

This follows from the Fourier transform of an exponential decay

\[ s(t) = e^{-2\pi |t|/\tau}, \]

which is Lorentzian\(^1\)

\[ \hat{s}(f) = \frac{1}{\pi} \frac{\tau^{-1}}{f^2 + \tau^{-2}}. \]

\(^1\)For a proof of this statement, see:
http://mathworld.wolfram.com/FourierTransformExponentialFunction.html
$\tau$ is inversely related to the width of the Lorentzian.

Let us fit only the real part. We shall determine the decay constant.

```matlab
figure; plot(real(signoise));
fid = real(signoise);
```

Let us also create a time axis. For example, suppose this FID was sampled over a 100ms time window. We create a vector containing 207 time values from 0 to 100 ms (207 is the number of points in the FID):

```matlab
taxis = linspace(0, 100, 207);
```

To launch the curve fitting GUI type `cftool`. In the drop-down box “X data” select the dataset “taxis” and in the box “Y data” select “fid”. `cftool` immediately attempts to fit a first degree polynomial (straight line):

Obviously, the fit is not a very good one, as evidenced by the low $R^2$ value of 0.4698. The other sign of the bad fit is the very large error bars on the fitted parameters $p_1$ and $p_2$.

Now let’s plot the residuals by hitting the “Residuals plot” button of the `cftool` window (second icon in the top left corner of the window). As you can see the residuals indicate that our model inadequately describes this
Let’s now select an appropriate model to fit this data. You can select from the drop-down menu “Exponential”, which will fit the data to a function of the form $a \times \exp(b \times x)$. The result is much better ($R^2$ value is 0.9876):

The residuals look fine, i.e. do not show any obvious structure. Let us now check whether or not adding a “baseline” to the fitting function leads to improvement. From the drop-down box, choose “Custom Equation” and in
the box below enter the function $a \cdot \exp(-b \cdot x) + c$. On my computer, I get the following error message:

```
Inf computed by model function, fitting cannot continue.
Try using or tightening upper and lower bounds on coefficients.
```

MATLAB is telling us that it cannot fit the data properly - some parameters go to infinity. We need to provide upper and lower bounds on the parameters. Click on the box “Fit Options”. You will see a new dialog box which provides several options for the fit. Since we know that the parameters “b” and “a” should be positive numbers, we enter 0 as the lower bound.

![Fit Options dialog box](image.png)

You can also check that in the drop-down box for “Algorithm” there are two options: **Trust-Region** and **Levenberg-Marquardt**.

These lower bounds are enough to cause the fit to converge:

---

2It’s possible that you don’t get this error. It depends on the particular version of MATLAB you are using; the default parameters may differ.
We note that the $R^2$ value (0.9877) only shows marginal improvement. Thus, the “baseline” parameter was superfluous in this case.

Finally, you can check that the residuals follow a Gaussian distribution. To do this, we will follow several steps. First, choose “Save to Workspace” from the “Fit” menu and click OK:

You can check that MATLAB has created new variables in the workspace. Clicking OK results in:

Variables have been created in the current workspace.

Then check by typing:

```
whos
```

and see:

<table>
<thead>
<tr>
<th>Name</th>
<th>Size</th>
<th>Bytes</th>
<th>Class</th>
<th>Attributes</th>
</tr>
</thead>
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<td>207x1</td>
<td>3312</td>
<td>double</td>
<td>complex</td>
</tr>
<tr>
<td>fid</td>
<td>1x207</td>
<td>1656</td>
<td>double</td>
<td></td>
</tr>
<tr>
<td>fittedmodel</td>
<td>1x1</td>
<td>961</td>
<td>cfit</td>
<td></td>
</tr>
</tbody>
</table>
One of these variables is the summary of results from the fit:

```matlab
fittedmodel
```

```matlab
goodness
```

```matlab
output
```
10. Learning to Program Optimization Algorithms in MATLAB

```
firstorderopt: 2.0225e-004
iterations: 6
funcCount: 28
cgiterations: 0
algorithm: 'trust-region-reflective'
message: [1x86 char]
```

To get the residuals, we use the notation `output.residuals`

```
figure;plot(taxis,output.residuals,'o-');xaxis('time');title('residuals');
```

We can also plot a histogram of the residuals by typing:

```
figure;hist(output.residuals,15);
```

The result is:

10.3. Steepest Descent Algorithm – implementation from scratch

We will now implement the steepest descent algorithm without the help of any built-in MATLAB fitting routines. This implementation could be ported to a low-level programming language such as FORTRAN or C with minimal effort. The update rule for this iterative algorithm is:

\[
\tilde{\beta}_{s+1} = \tilde{\beta}_s - \lambda \nabla \chi^2(\tilde{\beta}_s)
\]
where $\lambda$ is a parameter. If $\lambda$ is too large, the solutions will oscillate, and if it is too small, the algorithm will take too long to converge. $\chi^2$ is defined as the sum of square errors normalized to the square of errors in each of the $N$ data points $\{y_i\}$:

$$\chi^2(\vec{\beta}) = \sum_{i=1}^{n} \frac{(y_i - y(x_i; \vec{\beta}))^2}{\sigma_i^2}$$

where $y(x_i; \vec{\beta})$ is our fitting model.

10.3.1. **Fit to a straight line.** The fitting model is

$$y(x; A, B) = Ax + B.$$  

The gradient of $\chi^2$ has only two components:

$$(\nabla \chi^2)_A = \frac{\partial \chi^2}{\partial A} = \sum_{i=1}^{n} \frac{-2}{\sigma_i^2} (y_i - Ax_i - B)x_i$$

$$(\nabla \chi^2)_B = \frac{\partial \chi^2}{\partial B} = \sum_{i=1}^{n} \frac{-2}{\sigma_i^2} (y_i - Ax_i - B)$$

Let's create a MATLAB script by entering `edit fit3` whose content will be:

```matlab
1  tdata=linspace(0,10,100);  
2  a=20; b=2;  
3  ydata=a*tdata+b; % generate a fake data set  
4  sigma=10; % with noise added  
5  ydata=ydata+sigma*randn(size(ydata));  
6  figure;plot(tdata,ydata,'o-'); % visualize  
7  bs=[30 3]; % starting point  
8  sigmai=sigma*ones(size(ydata)); % assume error is constant for all ... measurements  
9  lambda=0.0001; % scaling parameter beta (if too large, solutions ... oscillate)  
10 for j=1:666,  
11    A=bs(1); % extract current value of A from the vector as  
12    B=bs(2); % extract value of B  
13    f=A*tdata+B; % fitting model function  
14    figure(1); hold off; plot(tdata,ydata,'bo-'); hold on;  
15    plot(tdata,f,'r');  
16    dx2da=sum((-2./(sigmai.^2)).*(ydata-A*tdata-B).*tdata);  
17    dx2db=sum((-2./(sigmai.^2)).*(ydata-A*tdata-B));  
18    gradchi=[dx2da dx2db]; % construct gradient vector  
19    bs = bs - lambda*gradchi; % bs+1=bs+h  
20    nchi2=sum((1./(sigmai.^2)).*(ydata-A*tdata-B).^2)/length(sigmai);  
21    fprintf(' iter=%d A=%f B=%f \chi^2=%f \n',j,bs(1),bs(2),nchi2);  
22 end
```
We obtain the following output

![Graph showing data points and fitted curve]

and fitting results:

```
iter=1  A=29.930332  B=2.989497  chi2=37.344001
iter=2  A=29.861141  B=2.979067  chi2=36.849307
...  
iter=665  A=19.937750  B=1.469166  chi2=1.118404
iter=666  A=19.937034  B=1.469032  chi2=1.118350
```

In this case, the steepest descent method performs reasonably well.

### 10.3.2. Nonlinear fit to an exponential decay.

Let’s consider the following model:

\[ y(x) = A \exp(-x/B) + C. \]

The \( \chi^2 \) function is

\[
\chi^2(\vec{\beta}) = \sum_{i=1}^{n} \frac{(y_i - A \exp(-x_i/B) - C)^2}{\sigma_i^2}
\]

where \( \vec{\beta} = (A, B, C) \).
10.3. Steepest Descent Algorithm – implementation from scratch

We can easily compute the gradient of $\chi^2$. It has three components, which are the partial derivatives of $\chi^2(A, B, C)$ with respect to each fitting parameter: $A$, $B$ and $C$:

$$(\nabla \chi^2)_A = \frac{\partial \chi^2}{\partial A} = \sum_{i=1}^{n} \frac{2}{\sigma_i^2} (y_i - A \exp(-x_i/B) - C) \exp(-x_i/B)$$

$$(\nabla \chi^2)_B = \frac{\partial \chi^2}{\partial B} = \sum_{i=1}^{n} \frac{2}{\sigma_i^2} (y_i - A \exp(-x_i/B) - C) \frac{A}{B^2} \exp(-x_i/B)$$

$$(\nabla \chi^2)_C = \frac{\partial \chi^2}{\partial C} = \sum_{i=1}^{n} \frac{2}{\sigma_i^2} (y_i - A \exp(-x_i/B) - C)$$

Let’s implement this algorithm. First we create a fake data set:

```matlab
1 tdata = linspace(0, 10, 100);
2 a = 500; b = 2; c = 40;
3 ydata = a * exp(-tdata/b) + c;
4 sigma = 30; % standard dev. of noise
5 ydata = ydata + sigma * randn(size(ydata));
6 figure; plot(tdata, ydata, 'o-');
```

Create a new MATLAB script file by typing `edit fit2` and copying the above code for generating our ‘fake’ data set. Then continue with the following commands:

```matlab
7 bs = [600 10 10]; % starting point
8 sigma = sigma*ones(size(ydata)); % assume error is constant for all ... measurements
9 lambda = 0.0001; % scaling parameter beta (if too large, solutions ... oscillate)
10 for j = 1:666,
11    A = bs(1); % extract current value of A from the vector as
12    B = bs(2); % extract value of B
13    C = bs(3); % extract C
14    f = A * exp(-tdata/B) + C; % fitting model function
15    figure(1); hold off; plot(tdata, ydata, 'bo-'); hold on;
16    plot(tdata, f, 'r');
17    dx2da = sum(-(2 ./ sigma.^2) .* (ydata - A * exp(-tdata/B) - C) .* exp(-tdata/B));
18    dx2db = sum(-(2 ./ sigma.^2) .* (ydata - A * exp(-tdata/B) - C) .* (A/B^2) .* exp(-tdata/B));
19    dx2dc = sum(-(2 ./ sigma.^2) .* (ydata - A * exp(-tdata/B) - C));
20    grad_chi = [dx2da dx2db dx2dc]; % construct gradient vector
21    bs = bs - lambda * grad_chi; % bs+1=bs+h
22    nchi2 = sum((1 ./ sigma.^2) .* (ydata - A * exp(-tdata/B) - C).^2) / length(sigma);
23    fprintf(' iter=%d A=%f B=%f C=%f chi2=%f ...
\r', j, bs(1), bs(2), bs(3), nchi2);
24 end
```
After almost 600 iterations we get the following output

![Graph showing data points and a fitted curve]

and fitting results are:

```
fit2

iter=1 A=599.996532 B=9.979192 C=9.996532 chi2=71.632804
iter=2 A=599.993071 B=9.958340 C=9.993071 chi2=71.444827
... iter=574 A=599.444376 B=1.627981 C=9.444376 chi2=2.979732
iter=575 A=599.444376 B=1.627981 C=9.444376 chi2=2.979732
```

The curve is fit, however, the fit results are not as good as with the FIT command. This is because steepest descent is not a very good algorithm compared to the one used by FIT. To find out which algorithm is used by FIT type:

```
help fitoptions
```

```
FITOPTIONS Create/modify a fit options object.
F = FITOPTIONS(LIBNAME) creates the fitoptions object F with the option parameters set to the default values for the library model LIBNAME. See CFLIBHELP for more information on LIBNAME.
```

and scroll down to the 'Algorithm' section:

```
Algorithm - Algorithm to be used in FIT
[{'Levenberg-Marquardt'} | 'Gauss-Newton' | 'Trust-Region']
```

Thus, the default algorithm is Levenberg-Marquardt which is inherently better than the steepest descent method. Other options are the Gauss-Newton and Trust-Region algorithms. If you’d like to try and compare any
of these algorithms, you can read the manual pages for FIT and FITOPTIONS to find out how to change the default method.

Something you may want to experiment with is trying different initial guesses. You will see that if the initial guess is too far from the actual values, this algorithm will fail miserably: in some cases it will get stuck in a local minimum and in other cases the solution may diverge. The second parameter you may want to experiment with is the value of $\eta$. Try a value 100 times larger or 100 times smaller.

10.4. Marquardt-Levenberg Algorithm

Marquardt-Levenberg update rule:

$$
\tilde{\beta}_{s+1} = \tilde{\beta}_s - (H_s + \lambda \text{diag}(H_s))^{-1} \nabla \chi^2(\tilde{\beta}_s)
$$

$H_s$ is proportional to the curvature of $\chi^2$, i.e. large steps are made in the direction of low curvature (flat terrain) and small steps in the direction with high curvature (steep incline). $H_s$ is called the curvature matrix.

The parameter $\lambda$ is adjusted at each iteration. We stop iterating when $\chi^2$ does not change appreciably. Here is a possible implementation of the Levenberg-Marquardt method:

- Pick initial guess for set of fitted parameters $\tilde{\beta}$.
- Compute $\chi^2(\tilde{\beta})$.
- Pick a modest value for $\lambda$, say $\lambda = 0.001$.
- $(*):$ Solve for $\delta\tilde{\beta} = (H_s + \lambda \text{diag}(H_s))^{-1} \nabla \chi^2(\tilde{\beta}_s)$ and evaluate $\chi^2(\tilde{\beta} + \delta\tilde{\beta})$.
- If $\chi^2(\tilde{\beta} + \delta\tilde{\beta}) \geq \chi^2(\tilde{\beta})$, we increase $\lambda$ by a factor of 10 and go back to $(*)$.
- If $\chi^2(\tilde{\beta} + \delta\tilde{\beta}) < \chi^2(\tilde{\beta})$, we decrease $\lambda$ by a factor of 10, update the trial solution $\tilde{\beta} \rightarrow \tilde{\beta} + \delta\tilde{\beta}$ and go back to $(*)$.
- Stopping criterion: changes in parameters that yield changes in $\chi^2$ by an amount $\ll 1$ are not statistically meaningful.
- When finished, use the Hessian to compute the estimated covariance matrix ($cov = 2H^{-1}$). The latter allows you to obtain errors in the fitted parameters or correlations among them.

10.4.1. Computer Implementation: Straight Line Fit. The update rule is:

$$
\tilde{\beta}_{s+1} = \tilde{\beta}_s - (H_s + \lambda \text{diag}(H_s))^{-1} \nabla \chi^2(\tilde{\beta}_s)
$$
where

\[ H_s \equiv H(\bar{\beta}_s) = \begin{bmatrix}
\frac{\partial^2 \chi^2}{\partial \beta_1^2} & \cdots & \frac{\partial^2 \chi^2}{\partial \beta_1 \partial \beta_p} \\
\vdots & \ddots & \vdots \\
\frac{\partial^2 \chi^2}{\partial \beta_p \partial \beta_1} & \cdots & \frac{\partial^2 \chi^2}{\partial \beta_p^2}
\end{bmatrix} \]

We have already seen how to compute the gradient vector \( \nabla \chi^2(\bar{\beta}_s) = (\frac{\partial \chi^2}{\partial \beta_1}, \ldots, \frac{\partial \chi^2}{\partial \beta_p})^T \).

In the case of the linear model, \( y(x) = Ax + B \).

The gradient of \( \chi^2 \)

\[ \chi^2(\bar{\beta}) = \sum_{i=1}^{n} \frac{(y_i - Ax_i - B)^2}{\sigma_i^2} \]

has the two components we previously found:

\[ (\nabla \chi^2)_A = \frac{\partial \chi^2}{\partial A} = \sum_{i=1}^{n} \frac{-2}{\sigma_i^2} (y_i - Ax_i - B)x_i \]

\[ (\nabla \chi^2)_B = \frac{\partial \chi^2}{\partial B} = \sum_{i=1}^{n} \frac{-2}{\sigma_i^2} (y_i - Ax_i - B) \]

whereas the Hessian matrix has 4 components:

\[ \frac{\partial^2 \chi^2}{\partial A^2} = \sum_{i=1}^{n} \frac{2x_i^2}{\sigma_i^2} \quad \quad \quad \quad \frac{\partial^2 \chi^2}{\partial A \partial B} = \sum_{i=1}^{n} \frac{2x_i}{\sigma_i^2} \]

\[ \frac{\partial^2 \chi^2}{\partial B \partial A} = \frac{2x_i}{\sigma_i^2} \quad \quad \quad \quad \frac{\partial^2 \chi^2}{\partial B^2} = \sum_{i=1}^{n} \frac{2}{\sigma_i^2} \]

The matrix \( H_s \) and its diagonal are:

\[ H_s = \begin{bmatrix}
\sum_{i=1}^{n} \frac{2x_i^2}{\sigma_i^2} & \sum_{i=1}^{n} \frac{2x_i}{\sigma_i^2} \\
\sum_{i=1}^{n} \frac{2x_i}{\sigma_i^2} & \sum_{i=1}^{n} \frac{2}{\sigma_i^2}
\end{bmatrix}, \quad \text{diag}[H_s] = \begin{bmatrix}
\sum_{i=1}^{n} \frac{2x_i^2}{\sigma_i^2} & 0 \\
0 & \sum_{i=1}^{n} \frac{2}{\sigma_i^2}
\end{bmatrix}. \]

The inverse of a \( 2 \times 2 \) matrix is easily obtained from the formula

\[ \begin{pmatrix} a & b \\ c & d \end{pmatrix}^{-1} = \frac{1}{ad - bc} \begin{pmatrix} d & -b \\ -c & a \end{pmatrix}. \]

For larger matrices, there are formulas to compute them. In the implementation of the algorithm on a computer, however, you will often resort to numerical techniques to obtain matrix inverses. In MATLAB you can use the \texttt{inv} command. There is also the pseudo-inverse \texttt{pinv} which works better when matrices are nearly singular.

For example, try
10.5. Curve Fitting by Simulated Annealing

>> A=[1 2 ; 3 4]

A =

   1  2
   3  4

inv(A)

ans =

  -2.0000    1.0000
   1.5000   -0.5000

and compare with the above formula. You should also check that this matrix inverse satisfies the definition of the inverse of a matrix: \( A^{-1}A = I = AA^{-1} \):

inv(A)*A

ans =

   1.0000    0.0000
   0.0000    1.0000

A*inv(A)

ans =

   1.0000    0.0000
   0.0000    1.0000

See Problem 81 for a complete implementation.

10.5. Curve Fitting by Simulated Annealing

Here is a possible MATLAB implementation of the Metropolis Monte-Carlo algorithm for curve fitting in the case of our previous example of an exponential decay function with noise.

1 % FIT_EXP_MC - Metropolis Monte-Carlo implementation of the curve
2 % fitting method
3
% create a 'fake' data set
tdata=linspace(0,10,100);
a=500; b=2; c=40;
% let's do an exponential decay
ydata=a*exp(-tdata/b)+c;
% and then add some gaussian random noise
sigma=30; % standard dev. of noise
ydata=ydata+sigma*randn(size(ydata));
figure;plot(tdata,ydata,'o-'); % visualize

as=[1 1 1]; % starting point, can pick randomly or not
sigmai=sigma*ones(size(ydata)); % assume error is constant for all
% measurements
T=1; % start with high temperature
for j=1:9999,
    A=as(1); % extract current value of A from the vector as
    B=as(2); % extract value of B
    C=as(3); % extract C
    f=A*exp(-tdata/B)+C; % fitting model function
    figure(1); hold off; plot(tdata,ydata,'bo-'); hold on;
    plot(tdata,f,'r'); drawnow;
    chi2=sum((1./(sigmai.^2)).*(ydata-A*exp(-tdata/B)-C).^2)/length(sigmai);
    \(\Delta E=0.05*randn([1 3])\); % random displacement vector
    asn=as+as.*\(\Delta \vec{b}\); % make random change in parameters
    An=asn(1); Bn=asn(2); Cn=asn(3);
    chi2n=sum((1./(sigmai.^2)).*(ydata-An*exp(-tdata/Bn)-Cn).^2)/length(sigmai);
    \(\Delta E=\Delta E\); % accept the move
    if (\(\Delta E<0\)), % accept the move
      as=asn;
    else, % else reject the move with probability P
      R=rand;
      P=exp(-\(\Delta E/T\));
      if (P > R),
        as=asn;
      end;
    end;
    T=T-0.01*T; % cooling schedule
    fprintf([' iter=%d A=%f B=%f C=%f chi2=%f \(\Delta E=%f\) T=%f 
', ...'
'|r'],j,as(1),as(2),as(3),chi2,DE,T);
end

A sample run proceeds as follows:

```
iter=1 A=332.450122 B=9.823360 C=9.608963 chi2=11.692912 deltaE=1.817395 T=0.990000
iter=2 A=332.450122 B=9.823360 C=9.608963 chi2=13.510307 deltaE=2.439683 T=0.970299
iter=3 A=317.456609 B=9.534128 C=9.311598 chi2=13.510307 deltaE=-1.658223 T=0.970299
```
10.7. Extrema Search by Simulated Annealing

iter=4  A=303.398269  B=9.540513  C=9.180057  chi2=11.852084
deltaE=-0.779770  T=0.960596

iter=665  A=521.159989  B=2.552933  C=4.891518  chi2=1.126674
deltaE=0.202851  T=0.001251

iter=666  A=521.159989  B=2.552933  C=4.891518  chi2=1.126674
deltaE=0.158424  T=0.001239

We have limited the number of iterations. However, in real implementations, a better criterion can be used, monitoring the value of $\chi^2$ to see if it changes significantly or not. I get something like this:

You should also try different initial guesses and check that the method is less sensitive to the choice of initial conditions as compared to the deterministic algorithms. It is easy to check, for example, that the steepest descent method does not work for the following choice of initial parameters, $[1 \ldots 1 1]$, because these parameters are too far from their true values.

10.6. Extrema Search by Newton Raphson Method

See Problem 81.

10.7. Extrema Search by Simulated Annealing

Can you explain what the code below does and how it works? What search (what type of extrema) is it performing? See also Problem 81.
Learning to Program Optimization Algorithms in MATLAB

```matlab
syms x
f=@(x) sin(x*0.01*pi)+0.25*sin(x*0.01*pi*5)+0.15*sin(x*0.01*pi*12);
figure(1); ezplot(f,[60 160]);

x=70; % start point
stepsize=5; % std dev
T=1; % start with high temperature
for j=1:600,
    xn=x+stepsize*randn(size(x)); % keep x within bounds!
    if (xn > 160), xn=160; end;
    if (xn < 60), xn=60; end;

    figure(2); clf; ezplot(f,[60 160]); hold on;
    plot(x,f(x),'ro'); drawnow; pause(0.1); % slow down!

    DE=f(xn)-f(x); % ∆E = final - initial
    if (DE < 0), x=xn; % accept move
    else, % else reject the move with probability P
        R=rand; P=exp(-DE/T);
        if (P > R), x=xn; end;
    end;
    T=T-0.01*T; % cooling schedule
    fprintf([' iter=%d f=%f ∆E=%f T=%f \n',j,f(x),DE,T);
end;
```

10.8. Problems

**Problem 79.** Create a fake data set for an exponential decay plus baseline, plus additive Gaussian noise:

```matlab
tdata=linspace(0,10,100);
a=500; b=2; c=40;
ydata=a*exp(-tdata/b)+c;
sigma=30; % standard dev. of noise
ydata=ydata+sigma*randn(size(ydata));
figure;plot(tdata,ydata,'o-');
```

(Addition of the noise with the `randn` command will introduce errors in the fitted parameters which you must determine.) Choose the model to be:

\[ y(x) = A \exp(-x/B) + C \]

Write a MATLAB program to implement the Gauss-Newton method (where the Hessian matrix is approximated by \(2J^TJ\), where \(J\) is the Jacobian matrix). Compare its performance to the Newton method. Does it converge faster or slower, how accurate are the results, and what do the errors in
the fitting parameters compare? Errors in the fitting parameters are obtained as the square root of the diagonal elements of the covariance matrix, $\text{cov}(\hat{\beta}, \hat{\beta}) = 2H^{-1}$, where $H$ is the Hessian matrix and $H^{-1}$ is its inverse.

Repeat the same exercise, except write MATLAB code to implement the Levenberg method. Compare method performance to Newton and Gauss-Newton methods.

Repeat the same exercise, except write MATLAB code to implement the Marquardt-Levenberg method. Compare method performance to Newton, Gauss-Newton and Levenberg methods.

Solution. These problems do not have unique solutions (everyone’s code can be different). However, the calculations for Jacobian matrix and other quantities are straightforward and follow procedures similar to what we did in class at the science learning center.

Problem 80. Write your own MATLAB code (as simple as possible) to implement the Newton Raphson method as a way:

(a) to find the zeros of the function

$$f(x) = x^3 - 9x^2 + 2x + 8$$

on the interval $[-2, 9]$.

Does the ability to find zeros depend on choice of initial condition? What algorithm do you think the MATLAB command \texttt{fzero} uses by default? Compare the performance of your code against the output of the \texttt{fzero} command.

(b) use Newton Raphson (and comparing with \texttt{fzero}) to find the point $x$ where $\exp(-\exp(-x)) = x$.

Solution. There are many possible solutions. Here is mine:

```matlab
1  % 1D Newton Raphson
2  clear
3  syms x
4  f=@(x) sin(x*0.01*pi)+0.25*sin(x*0.01*pi*5)+0.15*sin(x*0.01*pi*12);
5  f0=diff(f,x)
6  fpp=matlabFunction(diff(f0,x))
7  fpp=matlabFunction(diff(f0,x))
8  figure(1);
9  subplot(3,1,1),ezplot(f,[60 160]);
10 subplot(3,1,2),ezplot(f0,[60 160]);
11 subplot(3,1,3),ezplot(fpp,[60 160]);
12
13 xn=108; % start point
```
Problem 81. The following function of two independent variables \((x_1, x_2)\),
\[
f(x_1, x_2) = 20 + x_1^2 + x_2^2 - 10(\cos 2\pi x_1 + \cos 2\pi x_2),
\]
is plotted below:

As the plot shows, this function has many local minima. However, the function has just one global minimum, which occurs at the point [0 0] in the \(x_1-x_2\) plane, as indicated by the vertical line in the plot, where the value of the function is 0. At any local minimum other than [0 0], the value of the function is greater than 0. The farther the local minimum is from the origin, the larger the value of the function is at that point. As you can imagine, because this function has many local minima, it is difficult for standard, gradient-based methods to find the global minimum.

(a) Write MATLAB code to execute a gradient-based search (steepest descent) and show that the algorithm can at best find a local minimum near the starting point. Choose a starting point far from the global minimum, i.e. \(|x_1| > 3\) and \(|x_2| > 3\). On the other hand, show that if you pick a starting point near the origin (but not at the origin), the gradient search can find the global minimum. How far away from the origin can you pick the starting point, such that the search will still find the global minimum (i.e. such that beyond this radius, only a local minimum can be found)? In your algorithm, plot the function as a surface or contour plot, and indicate the position of the gradient-based search as function of time, so we can visualize the progress of the gradient search.
(b) Write simulated annealing code to search for the global minimum and show that the global minimum can always be located, regardless of the starting point.

**Solution.** (a) Here you are free to implement steepest descent, Newton, Gauss-Newton or Levenberg-Marquardt. We provide 2 examples; as you can see the code is fairly simple. For the Newton method:

```matlab
% Newton method in 2D
syms x1 x2
f = 8*(x1^2 + x2^2 - 10*cos(2*pi*x1)*cos(2*pi*x2)); % f
G=matlabFunction(gradient(f,[x1,x2])); % gradient of f
H=matlabFunction(hessian(f,[x1,x2])); % hessian of f
x=1*randn([2 1]); % initial position (x1,x2)
for j=1:50,
    x=x-pinv(H(x(1),x(2)))*G(x(1),x(2)); % Newton update rule
    figure(1); clf; ezcontour(f,[-4,4]); hold on;
    plot(x(1),x(2),'ro'); drawnow; pause(0.5);
    fprintf(' iter=%d f=%f x1=%f x2=%f
',j,f(x(1),x(2)),x(1),x(2));
end;
```

For the Levenberg-Marquardt method:

```matlab
% Marquardt-Levenberg method in 2D
syms x1 x2
f = 8*(x1^2 + x2^2 - 10*cos(2*pi*x1)*cos(2*pi*x2)); % f
G=matlabFunction(gradient(f,[x1,x2])); % gradient of f
H=matlabFunction(hessian(f,[x1,x2])); % hessian of f
x=1*randn([2 1]); % initial position (x1,x2)
lambda=0.001;
for j=1:50,
    hess=H(x(1),x(2));
    f_before=f(x(1),x(2));
    xn=x-pinv(hess+lambda*diag(diag(hess)))*G(x(1),x(2)); % update rule
    f_after=f(xn(1),xn(2));
    figure(1); clf; ezcontour(f,[-4,4]); hold on;
    plot(x(1),x(2),'ro'); drawnow; pause(0.5);
    fprintf(' iter=%d f=%f x1=%f x2=%f lambda=%f
',j,f(x(1),x(2)),x(1),x(2),lambda);
    if (f_after >= f_before), lambda=lambda*10;
    else lambda=lambda/10; x=xn;
end;
end;
```

(b) The following code converges in less than 4,000 steps. Initial temperature was T=12. Step size was 0.4. Cooling schedule was set to -0.1% per step.
10. Learning to Program Optimization Algorithms in MATLAB

```matlab
1 syms x1 x2
2 f = @(x1,x2) 20+x1^2+x2^2-10*(cos(2*pi*x1)+cos(2*pi*x2)); % f
3 x=1*randn([2 1]); % initial position (x1,x2)
4 T=12; % starting temperature
5 for j=1:20000,
6    figure(1); clf; ezcontour(f,[-4,4]); hold on;
7    plot(x(1),x(2),'ro'); drawnow; %pause(0.1);
8    dx=0.4*randn(size(x)); % make random move
9    f_before=f(x(1),x(2));
10   f_after=f(x(1)+dx(1),x(2)+dx(2));
11   dE=f_after-f_before; % change in energy
12   if (dE < 0), x=x+dx; % accept the move
13   else, % else reject move with probability P
14      R=rand; P=exp(-dE/T);
15      if (P > R), x=x+dx; end;
16   end;
17   T=T-0.001*T; % cooling schedule
18   fprintf(' iter=%d x1=%f x2=%f E=%f dE=%f T=%f 
', ... 
j,x(1),x(2),f(x(1),x(2)),dE,T);
19 end
```

Problem 82. Let $X$ and $Y$ be independent, identically distributed normal (Gaussian) random variables with mean 0 and variance 1. Write a computer program to generate random values for $X$ and $Y$. (Generate a very large number of pairs $X,Y$.) For each pair $(X,Y)$, compute the radius $R = \sqrt{X^2 + Y^2}$. Plot the probability density (i.e., histogram) for $R$. Obtain the resulting distribution.

Solution. $R$ follows a Rayleigh distribution. Also known as “Rice” or “Rician” distribution. Only 4 lines of code are needed to plot this distribution:

```matlab
1 X=randn([1 10000]);
2 Y=randn([1 10000]);
3 R=sqrt(X.^2+Y.^2);
4 figure;hist(R,50);
```
For those interested in a formal proof: consider the two-dimensional vector \((x, y)\) which has components that are Gaussian distributed, that is, \(p(x) = \frac{e^{-x^2/2\sigma^2}}{\sqrt{2\pi\sigma^2}}\), and the same for \(p(y)\). Denote with \(r\) the length \(r = \sqrt{x^2 + y^2}\). It is distributed as
\[
p(r|\sigma) = \frac{1}{2\pi\sigma^2} \int_{-\infty}^{\infty} dx \int_{-\infty}^{\infty} dy e^{-x^2/2\sigma^2} e^{-y^2/2\sigma^2} \delta(r - \sqrt{x^2 + y^2}).
\]
(The Dirac delta function is used to enforce the constraint that \(r\) must equal the length of the vector \((x, y)\).) By transforming to polar coordinates one has
\[
p(r|\sigma) = \frac{1}{2\pi\sigma^2} \int_0^{2\pi} d\phi \int_0^{\infty} d\rho \ \delta(r - \rho) \rho e^{-\rho^2/2\sigma^2} = \frac{r}{\sigma^2} e^{-r^2/2\sigma^2}
\]
which is the Rayleigh distribution.

**Problem 83.** Write a MATLAB program to fit an exponentially decaying signal. The signal \((ydata)\) can be simulated using the following sequence of commands (the code below generates the signal and adds noise):

```matlab
1 tdata=linspace(0,10,100);
2 a=500; b=2; c=40;
3 ydata=a*exp(-tdata/b)+c;
4 sigma=30; % standard dev. of noise
5 ydata=ydata+sigma*randn(size(ydata));
6 figure;plot(tdata,ydata,'o-');
```

You are allowed to use the following built-in MATLAB commands to fit the data: `fit`, `fminsearch`, `fminbnd`, `fminunc`, `fmincon`, `lsqnonlin`, ... `lsqcurvefit`. You are not allowed to use graphical user interfaces such as `cftool` - although it might be a useful exercise to experiment with it, as a way to double-check the performance of your method.

**Problem 84.** A number \(x\) is called a zero of a function \(f\) if and only if \(f(x) = 0\). One way to find zeros is to sketch the graph of the function. For
example, consider the function \( f(x) = x^3 - 4x^2 - 11x + 30 \). The following code could be used for this purpose:

```matlab
1 f=@(x) x.^3-4*x.^2-11*x+30;
2 x=linspace(-10,10,200);
3 y=f(x);
4 plot(x,y);
5 axis([-10,10,-100,100]);
6 grid on;
```

As you can see, the zeros are evident from the plot. You can also double-check that the zeros are indeed zeros. Suppose that the zeros are \(-3, 2, 5\). Then type:

```matlab
1 x=[-3,2,5];
2 f(x)
```

(a) In MATLAB check that \( x = -1 \) and \( x = 3 \) are zeros of the function \( f(x) = x^2 - 2x - 3 \) by plotting the function and identifying the zero crossings. (Plotting can be done as explained above, or an even easier way is to use the command \texttt{ezplot}.)

(b) Find the zeros of \( f(x) = x^4 - 29x^2 - 132 \).

(c) Use the MATLAB command \texttt{fzero} to find the zeros for the functions in (a) and (b). Does \texttt{fzero} give you a more accurate result than the graphical method?

(d) Use the command \texttt{fzero} to solve the following equation \( 5 - 2x = e^{-0.25x} \) for \( x \).

(e) Use the \texttt{fzero} command to find the zeros of the functions below on the given domain:

\[
\begin{align*}
  f(x) &= 9 - 4x - x^2 \quad \text{on } [-7, 3] \\
  f(x) &= 2x^2 - x - 8 \quad \text{on } [-3, 4] \\
  f(x) &= x^3 - 9x^2 + 2x + 8 \quad \text{on } [-2, 9]
\end{align*}
\]

(You will need to read the instructions on how to use the \texttt{fzero} command on an interval.)

**Problem 85.** We previously found that for an unrestricted random walk, the statistics of the random walker after \( n \) steps (histogram of \( S_n \)) follow a Gaussian distribution whose variance scales with \( n \). Write two MATLAB codes to simulate 1) a random walk with and 2) without boundary. For boundary: Place the boundary at some distance from the origin, but close enough, so that the particle hits the boundary. The boundary is reflective: when the particle hits the boundary, instead of crossing it, it is reflected.
back to the direction of origin. Plot the statistics of this random walk. The resulting distribution is clearly non-Gaussian. What distribution do you find?

**Problem 86.** Explain what this MATLAB code does. (pinv is nearly the same as inv.) What type of problem is it solving?

```matlab
1 syms x1 x2
2 f = @(x1,x2) 20+x1.^2+x2.^2-10*(cos(2*pi*x1)+cos(2*pi*x2)); % f
3 G=matlabFunction(gradient(f,[x1,x2])); % gradient of f
4 H=matlabFunction(hessian(f,[x1,x2])); % hessian of f
5 x=1*randn([2 1]); % initial position (x1,x2)
6 for j=1:50, % # of iterations
7     x=x-pinv(H(x(1),x(2)))*G(x(1),x(2));
8     figure(1); clf; ezcontour(f,[-4,4]); hold on; plot
9        plot(x(1),x(2),'ro'); drawnow;
10    fprintf(' iter=%d f=%f x1=%f x2=%f
',j,f(x(1),x(2)),x(1),x(2));
11 end;
```

**Solution.** This code implements the Newton method to iteratively find a local extremum of the function $f$. ■
Bibliography