Inferring the success parameter p of a binomial model from small samples affected by background

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Abstract

The problem of inferring the binomial parameter p from x *successes* obtained in n *trials* is reviewed and extended to take into account the presence of background, that can affect the data in two ways: *a*) fake successes are due to a background modeled as a Poisson process of known intensity; *b*) fake trials are due to a background modeled as a Poisson process of known intensity, each trial being characterized by a known success probability p_b.

1 Introduction

An important class of experiments consists in counting 'objects'. In fact, we are often interested in measuring their *density* in time, space, or both (here 'density' stands for a general term, that in the domain of time is equivalent to 'rate') or the *proportion* of those objects that have a certain character in common. For example, particle physicists might be interested in cross sections and branching ratios, astronomers in density of galaxies in a region of the sky or in the ratio of galaxies exhibiting some special features.

A well known problem in counting experiments is that we are rarely in the ideal situation of being able to count individually and at a given time all the objects of interest. More often we have to rely an a sample of them. Other problems that occur in real environments, especially in frontier research, are detector inefficiency and presence of background: sometimes we lose objects in counting; other times we might be confused by other objects that do not belong to the classes we are looking for, though they are observationally indistinguishable from the objects of interest.

We focus here on the effect of background in measurements of proportions. For a extensive treatment of the effect of background on rates, i.e. measuring the intensity of a Poisson process in presence of background, see Ref. [1], as well as chapters 7 and 13 of Ref. [2].

The paper is structured as follows. In section 2 we introduce the 'direct' and 'inverse' probabilistic problems related to the binomial distribution and the two cases of background that will be considered. In section 3 we go through the standard text-book case in which background is absent, but we discuss also, in some depth, the issue of how prior knowledge does or does not influence the probabilistic conclusions. Then, in the following two sections we come to the specific issue of this paper, and finally the paper ends with the customary short conclusions.

2 The binomial distribution and its inverse problem

An important class of counting experiments can be modeled as independent Bernoulli trials. In each trial we believe that a *success* will occur with probability p, and a *failure* with probability $q = 1 - p$. If we consider n independent trials, all with the same probability p, we might be interested in the total number of successes, independently of their order. The total number of successes X can range between 0 and n, and our belief on the outcome $X = x$ can be evaluated from the probability of each success and some combinatorics. The result is the well known *binomial* distribution, hereafter indicated with $\mathcal{B}_{n,p}$:

$$f(x \,|\, \mathcal{B}_{n,p}) = \frac{n!}{(n-x)!\,x!}\,p^x\,(1-p)^{n-x}\,, \qquad \left\{ \begin{array}{l} n = 1, 2, \ldots, \infty \\ 0 \le p \le 1 \\ x = 0, 1, \ldots, n \end{array} \right. \,, \qquad (1)$$

having *expected value* and *standard deviation*

$$\begin{aligned} \mathrm{E}(X) &= n\,p & (2) \\ \sigma(x) &= \sqrt{n\,p\,(1-p)}\,. & (3) \end{aligned}$$

We associate the formal quantities expected value and standard deviation to the concepts of (probabilistic) *prevision* and *standard uncertainty.*

The binomial distribution describes what is sometimes called a *direct probability* problem, i.e. calculate the probability of the experimental outcome x (the *effect*) given n and an assumed value of p. The *inverse* problem is what concerns mostly scientists: *infer p given n and x.* In probabilistic terms, we are interested in $f(p \,|\, n, x)$. Probability inversions are performed, within probability theory, using Bayes theorem, that in this case reads

$$f(p \,|\, x, n, \mathcal{B}) \;\propto\; f(x \,|\, \mathcal{B}_{n,p}) \cdot f_\circ(p) \qquad (4)$$

where $f_\circ(p)$ is the *prior*, $f(p\,|\,x,n,\mathcal{B})$ the *posterior* (or *final*) and $f(x\,|\,\mathcal{B}_{n,p})$ the *likelihood*. The proportionality factor is calculated from normalization. [Note the use of $f(\cdot)$ for the several probability functions as well as probability density functions (pdf), also within the same formula.] The solution of Eq. (4), related to the names of Bayes and Laplace, is presently a kind of first text book exercise in the so called Bayesian inference (see e.g. Ref. [2, 3]). The issue of priors in this kind of problems will be discussed in detail in Sec. 3.1, especially for the critical cases of $x=0$ and $x=n$.

The problem can be complicated by the presence of background. This is the main subject of this paper, and we shall focus on two kinds of background.

a) **Background can only affect x.** Think, for example, of a person shooting n times on a target, and counting, at the end, the numbers of scores x in order to evaluate his efficiency. If somebody else fires by mistake at random on his target, the number x will be affected by background. The same situation can happen in measuring efficiencies in those situations (for example due to high rate or loose timing) in which the time correlation between the equivalents of 'shooting' and 'scoring' cannot be done on a event by event basis (think, for example, to neutron or photon detectors).

The problem will be solved assuming that the background is described by a Poisson process of well known intensity r_b, that corresponds to a well known expected value λ_b of the resulting Poisson distribution (in the time domain $\lambda_b = r_b\cdot T$, where T is measuring time). In other words, the observed x is the sum of two contributions: x_s due to the *signal*, binomially distributed with $\mathcal{B}_{n,p}$, plus x_b due to background, Poisson distributed with parameter λ_b, indicated by $\mathcal{P}_{\lambda_b}$.

For large numbers (and still relatively low background) the problem is easy to solve: we subtract the expected number of background and calculate the proportion $\hat{p} = (x-\lambda_b)/n$. For small numbers, the 'estimator' $\hat{p}$ can become smaller than 0 or larger then 1. And, even if $\hat{p}$ comes out in the correct range, it is still affected by large uncertainty. Therefore we have to go through a rigorous probability inversion, that in this case is given by

$$f(p\,|\,n,x,\lambda_b) \quad\propto\quad f(x=x_s+x_b\,|\,n,p,\lambda_b)\cdot f_\circ(p)\,, \qquad (5)$$

where we have written explicitly in the likelihood that x is due to the sum of two (individually unobservable!) contributions x_s and x_b (hereafter the subscripts s and b stand for *signal* and *background*.)

b) **The background can show up, at random, as independent 'fake' trials, all with the same $\mathbf{p}_b$ of producing successes**. An example, that has indeed prompted this paper, is that of the measuring the proportion of blue galaxies in a small region of sky where there are galaxies belonging to a cluster, as well as background galaxies, the average proportion of blue galaxies of which is well known. In this case both n and x have two contributions:

$$\begin{aligned} n &= n_s + n_b & (6) \\ x &= x_s + x_b & (7) \end{aligned}$$

with

$$\begin{aligned} n_b &\sim \mathcal{P}_{\lambda_b} & (8) \\ x_b &\sim \mathcal{B}n_b, p_b & (9) \\ x_s &\sim \mathcal{B}n_s, p_s\,, & (10) \end{aligned}$$

where '$\sim$' stands for 'follows a given distribution'.

Again, the trivial large number (and not too large background) solution is the proportion of background subtracted numbers, $\hat{p} = (x - p_b\,\lambda_b)/(n - \lambda_b)$. But in the most general case we need to infer p from

$$f(p_s\,|\,n, x, \lambda_b, p_b) \quad \propto \quad f(x = x_s + x_b\,|\,n = n_s + n_b, p_b, \lambda_b) \cdot f_\circ(p)\,. \tag{11}$$

We might be also interested also to other questions, like e.g. how many of the n object are due to the signal, i.e.

$$f(n_s\,|\,n, x, \lambda_b, p_b)\,.$$

Indeed, the general problem lies in the joint inference

$$f(n_s, p_s\,|\,n, x, \lambda_b, p_b),$$

from which we can get other information, like the conditional distribution of p_s for any given number of events attributed to signal:

$$f(p_s\,|\,n, n_s, x, \lambda_b, p_b)\,.$$

Finally, we may also be interested in the rate r_s of the signal objects, responsible of the n_s signal objects in the sample (or, equivalently, to the Poisson distribution parameter λ_s):

$$f(\lambda_s\,|\,n, x, \lambda_b, p_b)\,.$$

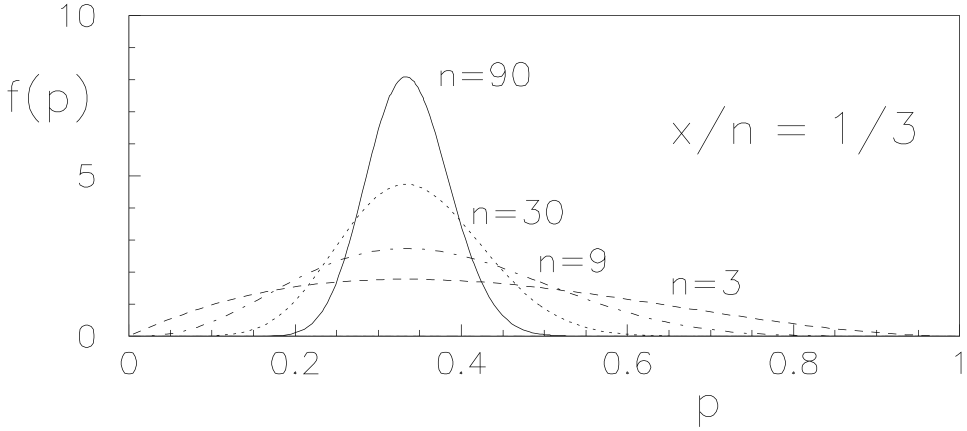


Figure 1: Probability density function of the binomial parameter p, having observed x successes in n trials.[2]

3 Inferring p in absence of background

The solution of Eq.(4) depends, at least in principle, on the assumption on the prior $f_\circ(x)$. Taking a flat prior between 0 and 1, that models our *indifference* on the possible values of p *before* we take into account the result of the experiment in which x successes were observed in n trials, we get (see e.g. [2]):

$$f(p\,|\,x, n, \mathcal{B}) = \frac{(n+1)!}{x!\,(n-x)!}\,p^x\,(1-p)^{n-x}\,, \tag{12}$$

some examples of which are shown in Fig. 1. Expected value, mode (the value of p for which $f(p)$ has the maximum) and variance of this distribution are:

$$\begin{aligned} \mathrm{E}(p) &= \frac{x+1}{n+2} && (13)\\ \mathrm{mode}(p) = p_m &= x/n && (14)\\ \sigma^2(p) = \mathrm{Var}(p) &= \frac{(x+1)(n-x+1)}{(n+3)(n+2)^2} && (15)\\ &= \mathrm{E}(p)\,(1-\mathrm{E}(p))\,\frac{1}{n+3}\,. && (16) \end{aligned}$$

Eq. (13) is known as “recursive Laplace formula”, or “Laplace's rule of succession”. Not that there is no magic if the formula gives a sensible result even for the extreme cases $x = 0$ and $x = n$ for all values of n (even if $n = 0$!).

It is just a consequence of the prior: in absence of new information, we get out what we put in!

From Fig. 1 we can see that for large numbers (and with x far from 0 and from n) $f(p)$ tends to a Gaussian. This is just the reflex of the limit to Gaussian of the binomial. In this large numbers limit $\mathrm{E}(p) \approx p_m = x/n$ and $\sigma(p) \approx \sqrt{x/n\,(1-x/n)/n}$.

3.1 Meaning and role of the prior: many data limit versus frontier type measurements

One might worry about the role of the prior. Indeed, in some special cases of importance *frontier type* measurement one *has* to. However, in most *routine* cases, the prior just plays the role of a *logical tool to allow probability inversion*, but it is in fact absorbed in the normalization constant. (See extensive discussions in Ref. [2] and references therein.)

In order to see the effect of the prior, let us model it in a easy and powerful way using a *beta* distribution, a very flexible tool to describe many situations of prior knowledge about a variable defined in the interval between 0 and 1 (see Fig. 2). The beta distribution is the *conjugate prior* of the binomial distribution, i.e. prior and posterior belong to the same function family, with parameters updated by the data via the likelihood. In fact, a generic beta distribution in function of the variable p is given by

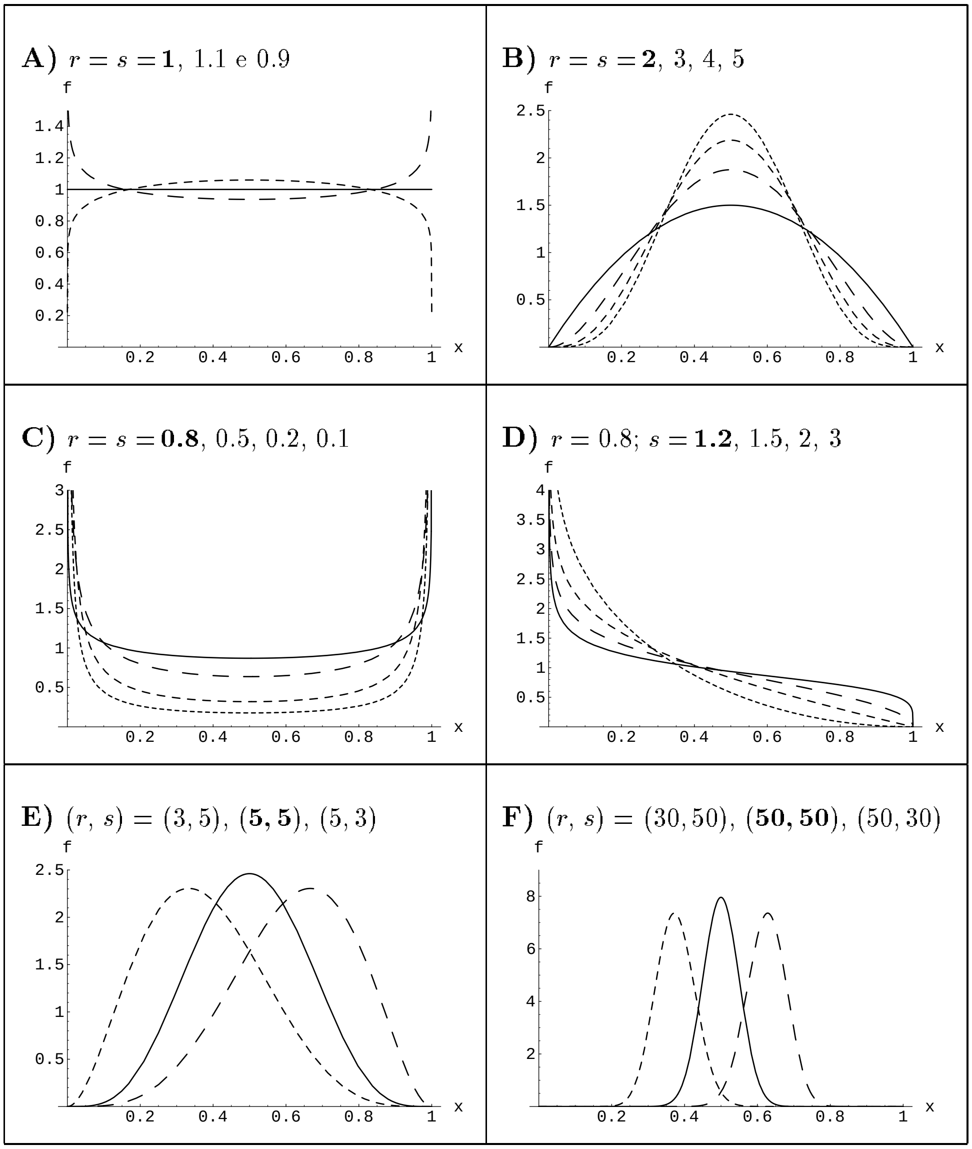
$$f(p\,|\,\mathrm{Beta}(r,s)) = \frac{1}{\beta(r,s)} p^{r-1}(1-p)^{s-1} \qquad \left\{ \begin{array}{l} r,\, s>0 \\ 0 \le p \le 1\,. \end{array} \right. \tag{17}$$

The denominator is just for normalization and, indeed, the integral $\beta(r,s) = \int_0^1 p^{r-1}(1-p)^{s-1}\,\mathrm{d}p$ defines the special function beta that names the distribution. We immediately recognize Eq. (12) as a beta distribution of parameters $r = x+1$ and $s = n-x+1$ [and the fact that $\beta(r,s)$ is equal to $(r-1)!(s-1)!/(s+r-1)!$ for integer arguments].

For a generic beta we get the following posterior (neglecting the irrelevant normalization factor):

$$\begin{aligned} f(p\,|\,n,x,\mathrm{Beta}(r,s)) &\propto \left[p^x(1-p)^{n-x}\right] \times \left[p^{r_i-1}(1-p)^{s_i-1}\right] &(18)\\ &\propto p^{x+r_i-1}(1-p)^{n-x+s_i-1}\,, &(19)\end{aligned}$$

where the subscript i stands for *initial*, synonym of prior. We can then see that the final distribution is still a beta with parameters $r_f = r_i + x$ and $s_f = s_i + (n-x)$: the first parameter is updated by the number of successes, the second parameter by the number of failures.



A) $r = s = \mathbf{1}$, 1.1 e 0.9

B) $r = s = \mathbf{2}$, 3, 4, 5

C) $r = s = \mathbf{0.8}$, 0.5, 0.2, 0.1

D) $r = 0.8$; $s = \mathbf{1.2}$, 1.5, 2, 3

E) $(r,\, s) = (3,5)$, $(\mathbf{5,5})$, $(5,3)$

F) $(r,\, s) = (30,50)$, $(\mathbf{50,50})$, $(50,30)$

Figure 2: Examples of Beta distributions for some values of r and s [2]. The parameters in bold refer to continuous curves.

Expected value, mode and variance of the generic beta of parameters r and s are:

$$\begin{aligned}
\mathrm{E}(X) &= \frac{r}{r+s} && (20)\\
\mathrm{mode}(X) &= (r-1)/(r+s-2) \qquad [r>1 \text{ and } s>1] && (21)\\
\mathrm{Var}(X) &= \frac{rs}{(r+s+1)\,(r+s)^2} \qquad [r+s>1]\,. && (22)
\end{aligned}$$

Then we can use these formulae for the beta posterior of parameters r_f and s_f.

The use of the conjugate prior in this problem demonstrates in a clear way how the inference becomes progressively independent from the prior information in the limit of a large amount of data: this happens when both $x \gg r_i$ and $n-x \gg s_i$. In this limit we get the same result we would get from a flat prior ($r_i = s_i = 1$, see Fig. 2). For this reason in standard 'routine' situation, we can quietly and safely take a flat prior.

Instead, the treatment needs much more care in situations typical of 'frontier research': small numbers, and often with no single 'successes'. Let us consider the latter case and let us assume a naïve flat prior, that it is considered to represent 'indifference' of the parameter p between 0 and 1. From Eq. (12) we get

$$f(p\,|\,x=0,n,\mathcal{B},\mathrm{Beta}(1,1)) \;=\; (n+1)\,(1-p)^n\,. \qquad (23)$$

(The prior has been written explicitly among the conditions of the posterior.) Some examples are given in Fig. (3). As n increases, p is more and more constrained in proximity of 0. In these cases we are used to give *upper limits* at a certain level of confidence. The natural meaning that we give to this expression is that we are such and such percent confident that p is below the reported upper limit. In the Bayesian approach this is is straightforward, for confidence and probability are synonyms. For example, if we want to give the limit that makes us 95% sure that p is below it, i.e. $P(p \le p_{u_{0.95}}) = 0.95$, then we have to calculate the value $p_{u_{0.95}}$ such that the cumulative function $F(p_{u_{0.95}})$ is equal to 0.95:

$$\begin{aligned}
F(p_{u_{0.95}}\,|\,x=0,n,\mathcal{B},\mathrm{Beta}(1,1)) &= \int_0^{p_{u_{0.95}}} f(p)\,\mathrm{d}p && (24)\\
&= 1-(1-p_u)^n = 0.95\,, && (25)
\end{aligned}$$

that yields

$$p_{u_{0.95}} \;=\; 1-\sqrt[n+1]{0.05}\,. \qquad (26)$$

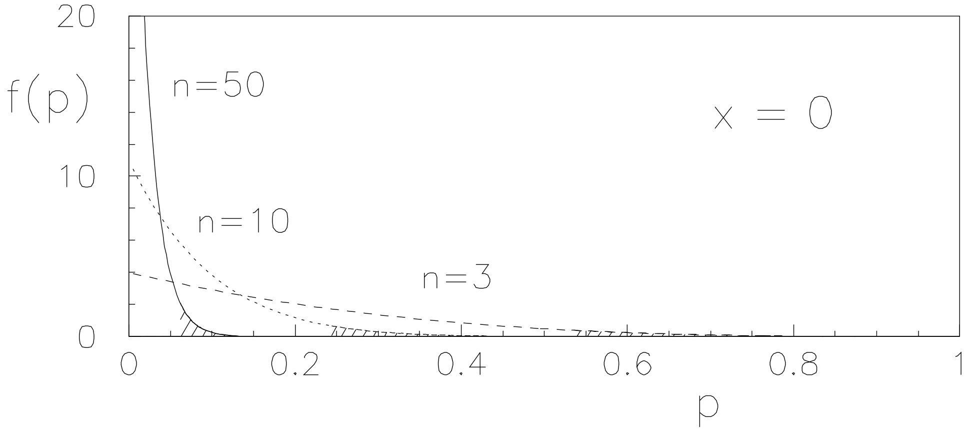


Figure 3: Probability density function of the binomial parameter p, having observed no successes in n trials.[2]

For the three examples given in Fig. 3, with n = 3, 10 and 50, we have $p_{u_{0.95}}$ = 0.53, 0.24 and 0.057, respectively. These results are in order, as long the flat prior reflected our expectations about p, that it could be about equally likely in any sub-interval of fixed width in the interval between 0 and 1 (and, for example, we believe that it is equally likely below 0.5 and above 0.5).

However, this is often not the case in frontier research. Perhaps we were looking for a very rare process, with a very small p. Therefore, having done only 50 trials, we cannot say to be 95% sure that p is below 0.057. In fact, by logic, the previous statement implies that we are 5% sure that p is above 0.057, and this might seem too much for the scientist expert of the phenomenology under study. (Never ask mathematicians about priors! Ask yourselves and the colleagues you believe are the most knowledgeable experts of what you are studying.) In general I suggest to make the exercise of calculating a 50% upper or lower limit, i.e. the value that divides the possible values in two equiprobable regions: we are as confident that p is above as it is below $p_{u_{0.5}}$. For $n = 50$ we have $p_{u_{0.5}} = 0.013$. If a physicist was looking for a rare process, he/she would be highly embarrassed to report to be 50% confident that p is above 0.013. But he/should be equally embarrassed to report to be 95% confident that p is below 0.057, because both statements are logical consequence of the same result, that is Eq. (23). If this is the case, a better grounded prior is needed, instead of just a 'default' uniform. For example one might thing that several order of magnitudes in the small p

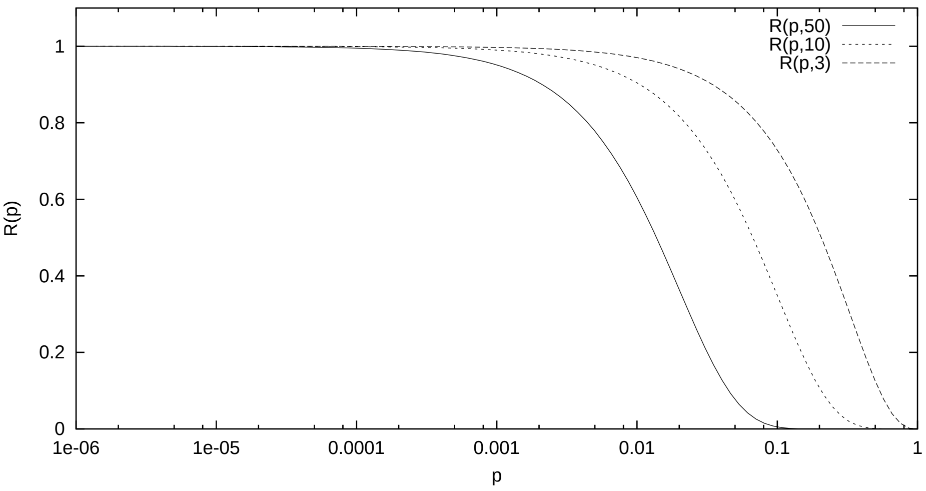


Figure 4: Rescaled likelihoods for $x = 0$ and some values of n

range are considered equally possible. This give rise to a prior that is uniform in $\ln p$ (within a range $\ln p_{min}$ and $\ln p_{max}$), equivalent to $f_\circ(p) \propto 1/p$ with lower and upper cut-off's.

Anyway, instead of playing blindly with mathematics, looking around for 'objective' priors, or priors that come from abstract arguments, it is important to understand at once the role of prior and likelihood. Priors are logically important to make a 'probably inversion' via the Bayes formula, and it is a matter of fact that no other route to probabilistic inference exists. The task of the likelihood is to modify our beliefs, distorting the pdf that models them. Let us plot the three likelihoods of the three cases of Fig. 3, *rescaled* to the asymptotic value $p \to 0$ (constant factors are irrelevant in likelihoods). It is preferable to plot them in a log scale along the abscissa to remember that several orders of magnitudes are involved (Fig. 4).

We see from the figure that in the high p region the beliefs expressed by the prior are strongly dumped. If we were convinced that p was in that region we have to dramatically review our beliefs. With the increasing number of trials, the region of 'excluded' values of $\log p$ increases too.

Instead, for very small values of p, the likelihood becomes flat, i.e. equal to the asymptotic value $p \to 0$. The region of flat likelihood represents the values of p for which the experiment loses sensitivity: if scientific motivated priors concentrate the probability mass in that region, then the experiment is irrelevant to change our convictions about p.

Formally the rescaled likelihood

$$\mathcal{R}(p;\, n,\, x=0) \;=\; \frac{f(x=0\,|\,n,\,p)}{f(x=0\,|\,n,\,p\to 0)}\,, \tag{27}$$

equal to $(1-p)^n$ in this case, is a functions that gives the Bayes factor of a generic p with respect to the reference point $p=0$ for which the experimental sensitivity is certainly lost. Using the Bayes formula, $\mathcal{R}(p;\, n,\, x=0)$ can rewritten as

$$\mathcal{R}(p;\, n,\, x=0) \;=\; \frac{f(p\,|\,n,\,x=0)}{f_\circ(p)} \bigg/ \frac{f(p=0\,|\,n,\,x=0)}{f_\circ(p=0)}\,, \tag{28}$$

to show that it can be interpreted as a *relative belief updating factor*, in the sense that it gives the updating factor for each value of p with respect to that at the asymptotic value $p\to 0$.

We see that this $\mathcal{R}$ function gives a way to report an upper limit that do not depend on prior: it can be any conventional value in the region of transition from $\mathcal{R}=1$ to $\mathcal{R}=0$. However, this limit cannot have a probabilistic meaning, because does not depend on prior. It is instead a *sensitivity bound*, roughly separating the excluded high p value from the the small p values about which the experiment has nothing to say.[1]

For further discussion about the role of prior in frontier research, applied to the Poisson process, see Ref. [1]. For examples of experimental results provided with the $\mathcal{R}$ function, see Refs. [4, 5, 6].

4 Poisson background on the observed number of 'successes'

Imagine now that the x successes might contains an unknown number of background events x_b, of which we only know their expected value λ_b, estimated somehow and about which we are quite sure (i.e. uncertainty about λ_b is initially neglected — it will be indicated at the end of the section how to handle it). We make the assumption that the background events come at random and are described by a Poisson process of intensity r_b, such that the Poisson parameter λ_b is equal to $r_b \times \Delta T$ in the domain of time, with ΔT the observation time. (But we could as well reason in other domains, like objects per unit of length, surface, volume, or solid angle. The density/intensity parameter r will have different dimensions depending on the context, while λ will always be dimensionless.)

[1] "*Wovon man nicht reden kann, darüber muss man schweigen*" (L. Wittgenstein).

The number of observed successes x has now two contributions:

$$\begin{aligned} x &= x_s + x_b & (29)\\ x_s &\sim \mathcal{B}_{n,p} & (30)\\ x_b &\sim \mathcal{P}_{\lambda_b}\,, & (31) \end{aligned}$$

In order to use Bayes theorem we need to calculate $f(x\,|\,n,\,p,\,\lambda_b)$, that is $f(x = x_s + x_b\,|\,\mathcal{B}_{n,p},\,\mathcal{P}_{\lambda_b})$, i.e. is the probability function of the sum of a binomial variable and a Poisson variable. The combined probability function is give by (see e.g. section 4.4 of Ref. [2]):

$$f(x\,|\,\mathcal{B}_{n,p},\,\mathcal{P}_{\lambda_b}) \;=\; \sum_{x_s,\,x_b} \delta_{x,\,x_s+x_b}\, f(x_s\,|\,\mathcal{B}_{n,p_s})\, f(x_b\,|\,\mathcal{P}_{\lambda_b}) \qquad (32)$$

where $\delta_{x,\,x_s+x_b}$ is the Kronecker delta that constrains the possible values of x_s and x_b in the sum (x_s and x_b run from 0 to the maximum allowed by the constrain). Note that we do not need to calculate this probability function for all x, but only for the number of actually observed successes.

The inferential result about p is finally given by

$$f(p\,|\,n,\,p,\,\lambda_b) \;\propto\; f(x\,|\,\mathcal{B}_{n,p},\,\mathcal{P}_{\lambda_b})\, f_0(p)\,. \qquad (33)$$

An example is shown in Fig. 5, for $n = 10$, $x = 7$ and an expected number of background events ranging between 0 and 10, as described in the figure caption. The upper plot of the figure is obtained by a uniform prior (priors are represented with dashed lines in this figure). As an exercise, let us also show in the lower plot of the figure the results obtained using a broad prior still centered at $p = 0.5$, but that excludes the extreme values 0 and 1, as it is often the case in practical cases. This kind of prior has been modeled here with a beta function of parameters $r_i = 2$ and $s_i = 2$.

For the cases of expected background different from zero we have also evaluated the $\mathcal{R}$ function, defined in analogy to Eq. (27) as $\mathcal{R}(p;\,n,\,x,\,\lambda_b) = f(x\,|\,n,\,p,\lambda_b)/f(x\,|\,n,\,p \to 0,\,\lambda_b)$. Note that, while Eq. (27) is only defined for $x \neq 0$, since a single observation makes $p = 0$ impossible, that limitation does not hold any longer in the case of not null expected background. In fact, it is important to remember that, as soon as we have background, there is some chance that all observed events are due to it (remember that a Poisson variable is defined for all non negative integers!). This is essentially the reasons why in this case the likelihoods tend to a positive value for $p \to 0$ (I like to call 'open' this kind of likelihoods [2]). As discussed above, the power of the data to update the believes on p is self-evident in a log-plot.

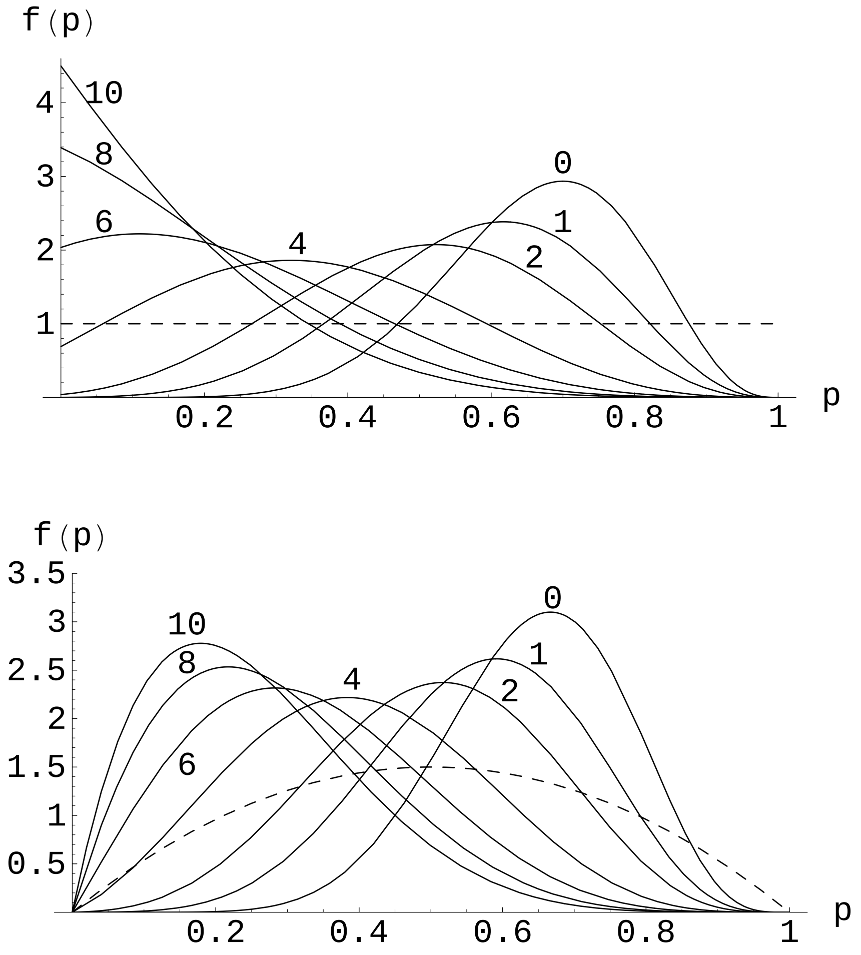


Figure 5: Inference of p for $n = 10$, $x = 7$, and several hypotheses of background (right to left curves for $\lambda_B = 0$, 1, 2, 4, 5, 6, 10) and two different priors (dashed lines), Beta$(1, 1)$ in the upper plot and Beta$(2, 2)$ in the lower plot (see text).

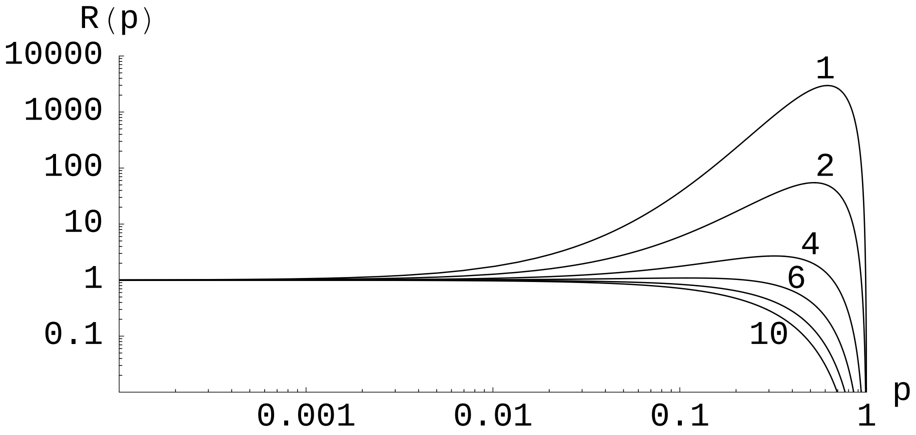


Figure 6: Relative believe updating factor of p for $n = 10$, $x = 7$ and several hypotheses of background: $\lambda_B = 1$, 2, 4, 6, 8, 10.

We seen in Fig. 6 that, essentially, the data do not provide any relevant information for values of p below 0.01.

Let us also see what happens when the prior concentrates our beliefs at small values of p, though in principle allowing all values of from 0 to 1. Such a prior can be modeled with a log-normal distribution of suitable parameters (-4 and 1), i.e. $f_0(p) = \exp\left[-(\log p + 4)^2)/2\right]/(\sqrt{2\,\pi}\,p)$, with an upper cut-off at $p = 1$ (the probability that such a distribution gives a value above 1 is $3.2\,10^{-5}$). Expected value and standard deviation of Lognormal(-4,1) are 0.03 and 0.04, respectively. The result is given in Fig. 7, where the prior is indicated with a dashed line.

We see that, with increasing expected background, the posteriors are essentially equal to the prior. Instead, in case of null background, ten trials are already sufficiently to dramatically change our prior beliefs. For example, initially there was 4.5% probability that p was above 0.1. Finally there is only 0.09% probability for p to be below 0.1.

The case of null background is also shown in Fig. 8, where the results of the three different priors are compared. We see that passing from a Beta(1, 1) to a Beta(2, 2), makes little change in the conclusion. Instead, a log-normal prior distribution peaked at low values of p changes quite a lot the shape of the distribution, but not really the substance of the result (expected value and standard deviation for the three cases are: 0.67, 0.13; 0.64, 0.12; 0.49, 0.16). Anyway, the prior does correctly its job and there should be no wonder that the final pdf drifts somehow to the left side, to

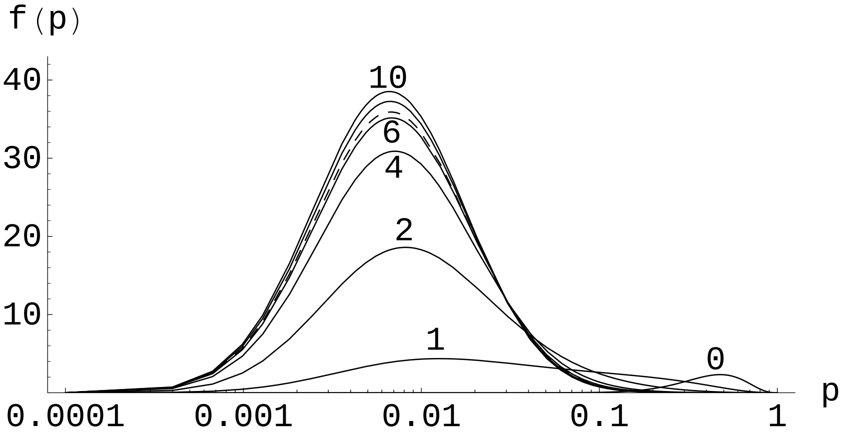


Figure 7: Inference of p for $n = 10$, $x = 7$, assuming a log-normal prior (dashed line) peaked at low p, and with several hypotheses of background ($\lambda_B = 0, 1, 2, 4, 6, 8, 10$).

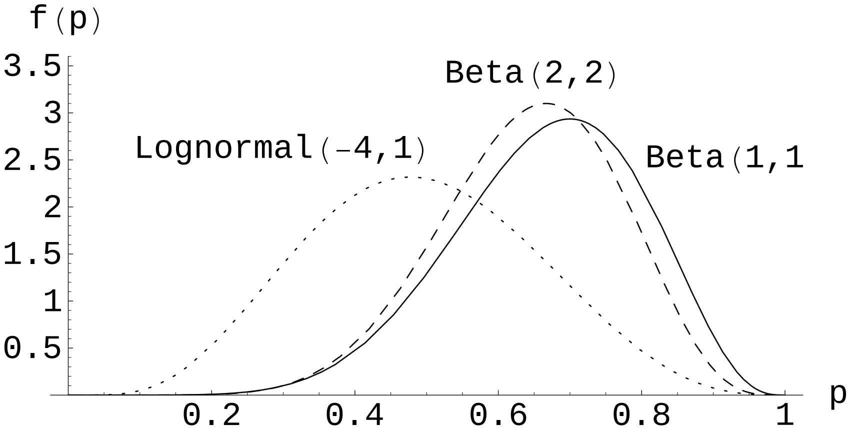


Figure 8: Inference of p for $n = 10$, $x = 7$ in absence of background, with three different priors.

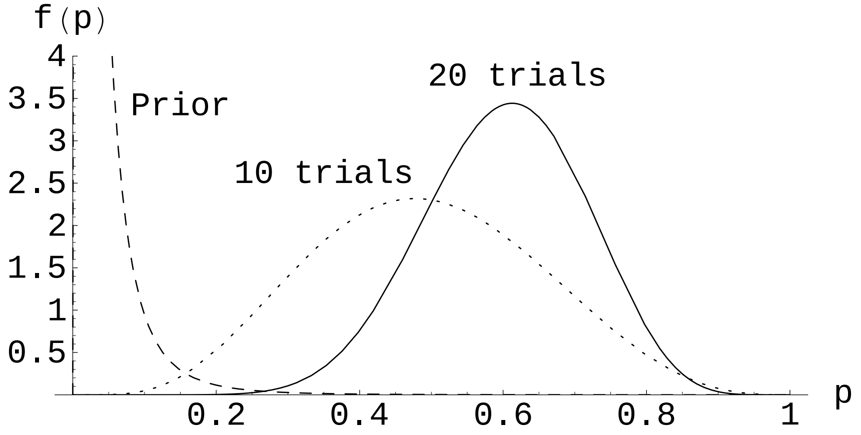


Figure 9: Sequential inference of p, starting from a prior peaked at low values, given two experiments, each with $n = 10$ and $x = 7$.

take into account a prior knowledge according to which 7 successes in 10 trials was really a 'surprising event'.

Those who share such a prior need more solid data to be convinced that p could be much larger than what they initially believed. Let make the exercise of looking at what happens if a second experiment gives exactly the same outcome ($x = 7$ with $n = 10$). The Bayes formula is applied sequentially, i.e. the posterior of the first inference become the prior of the second inference. That is equivalent to multiply the two priors (we assume conditional independence of the two observations). The results are given in Fig. 9. (By the way, the final result is equivalent to having observed 14 successes in 20 trials, as it should be — the correct updating property is one of the intrinsic nice features of the Bayesian approach).

4.1 Uncertainty on the expected background

In these examples we made the assumption that the expected number of background events is well known. If this is not the case, we can quantify our uncertainty about it by a pdf $f(\lambda_b)$, whose modeling depends on our best knowledge about λ_s. Taking account of this uncertainty in a probabilistic approach is rather simple, at least conceptually (calculations can be quite complicate, but this is a different question). In fact, applying probability theory we get:

$$f(p\,|\,x,\,n) \;=\; \int_0^\infty f(p\,|\,x,\,n,\,\lambda_b)\,f(\lambda_b)\,\mathrm{d}\lambda_b\,. \qquad (34)$$

We recognize in this formula that the pdf that takes into account all possible values of λ is a weighted average of all λ_b dependent pdf's, with a weight equal to $f(\lambda_b)$.

5 Poisson background on the observed number of 'trials' and of 'successes'

Let us know move to problem *b)* of the introduction. Again, we consider only the background parameters are well known, and refer to the previous subsection for treating their uncertainty. To summarize, that is what we assume to know with certainty:

n : the total observed numbers of 'objects', n_s of which are due to signal and n_b to background; but these two numbers are not directly observable and can only be inferred;

x : the total observed numbers of the 'objects' of the subclass of interest, sum of the unobservable x_s and x_b;

λ_b : the expected number of background objects;

p_b : the expected proportion of successes due to the background events.

As we discussed in the introduction, we are interested in inferring the number of signal objects n_s, as well as the parameter p_s of the 'signal'. We need then to build a likelihood that connects the observed numbers to all quantities we want to infer. Therefore we need to calculate the probability function $f(x\,|\,n,\; n_s,\; p_s,\; \lambda_b,\; p_b)$.

Let us first calculate the probability function $f(x\,|\,n_s,\; p_s\, n_b,\; p_b)$ that depends on the unobservable n_s and n_b. This is the probability function of the sum of two binomial variables:

$$f_{2\mathcal{B}}(x\,|\,n_s,\; p_s\, n_b,\; p_b) \;=\; \sum_{x_s,\, x_b} \delta_{x,\, x_s+x_b}\, f(x_s\,|\,\mathcal{B}_{n_s,p_s}) \cdot f(x_b\,|\,\mathcal{B}_{n_b,p_b})\,, \quad (35)$$

where x_s ranges between 0 and n_s, and x_b ranges between 0 and n_b. x can vary between 0 and $n_s + n_b$, has expected value $\mathrm{E}(x) = n_s\, p_s + n_b\, p_b$ and variance $\mathrm{Var}(x) = n_s\, p_s\, (1-p_s) + n_b\, p_b\, (1-p_b)$. As for Eq. (32), we need to evaluate Eq. (35) only for the observed number of successes. Contrary to the implicit convention within this paper to use the same symbol $f(\cdot)$ meaning different probability functions and pdf's, we name Eq. (35) $f_{2\mathcal{B}}$ for later convenience.

In order to obtain the general likelihood we need, two observations are in order:

- Since x depends from λ only via n_b, then $f(x\,|\,n_s,\,p_s\,n_b,\,p_b,\,\lambda_b)$ is equal to $f_{2\mathcal{B}}(x\,|\,n_s,\,p_s\,n_b,\,p_b)$.
- The likelihood that depends also on n can obtained from $f(x\,|\,n_s,\,p_s\,n_b,\,p_b,\,\lambda_b)$ by the following reasoning:
  - if $n = n_s + n_b$, then
  $$f(x\,|\,n,\,n_s,\,p_s\,n_b,\,p_b,\,\lambda_b) = f(x\,|\,n_s,\,p_s\,n_b,\,p_b,\,\lambda_b)\,;$$
  - else
  $$f(x\,|\,n,\,n_s,\,p_s\,n_b,\,p_b,\,\lambda_b) = 0\,.$$

It follows that

$$\begin{aligned} f(x\,|\,n,\,n_s,\,p_s,\,n_b,\,p_b,\,\lambda_b) &= f(x\,|\,n_s,\,p_s\,n_b,\,p_b,\,\lambda_b)\,\delta_{n,\,n_s+n_b} & (36)\\ &= f_{2\mathcal{B}}(x\,|\,n_s,\,p_s\,n_b,\,p_b)\,\delta_{n,\,n_s+n_b}\,. & (37)\end{aligned}$$

At this point we get rid of n_b in the conditions, taking account its possible values and their probabilities, given λ_b:

$$f(x\,|\,n,\,n_s,\,p_s,\,p_b,\,\lambda_b) = \sum_{n_b} f(x\,|\,n,\,n_s,\,p_s,\,n_b,\,p_b,\,\lambda_b)\,f(n_b\,|\,\mathcal{P}_{\lambda_b})\,, \qquad (38)$$

i.e.

$$f(x\,|\,n,\,n_s,\,p_s,\,p_b,\,\lambda_b) = \sum_{n_b} f_{2\mathcal{B}}(x\,|\,n_s,\,p_s\,n_b,\,p_b)\,f(n_b\,|\,\mathcal{P}_{\lambda_b})\,\delta_{n,\,n_s+n_b}\,, (39)$$

where n_b ranges between 0 and x, due to the $\delta_{n,\,n_s+n_b}$ condition. Finally, we can use Eq. (39) in Bayes theorem to infer n_s and p_s:

$$\begin{aligned} f(n_s,\,p_s\,|\,x,\,n,\,\lambda_b,\,p_b) &\propto f(x\,|\,n,\,n_s,\,p_s,\,p_b,\,\lambda_b)\,f_0(n_s,\,p_s) & (40)\\ f(p_s\,|\,x,\,n,\,\lambda_b,\,p_b) &= \sum_{n_s} f(n_s,\,p_s\,|\,x,\,n,\,\lambda_b,\,p_b) & (41)\\ f(n_s\,|\,x,\,n,\,\lambda_b,\,p_b) &= \int f(n_s,\,p_s\,|\,x,\,n,\,\lambda_b,\,p_b)\,\mathrm{d}p_s & (42)\\ f(p_s\,|\,x,\,n,\,n_s,\,\lambda_b,\,p_b) &= \frac{f(n_s,\,p_s\,|\,x,\,n,\,\lambda_b,\,p_b)}{f(n_s\,|\,x,\,n,\,\lambda_b,\,p_b)}\,. & (43)\end{aligned}$$

We give now some numerical examples. For simplicity (and because we are not thinking to a specific physical case) we take uniform priors, i.e. $f_0(n_s,\,p_s) = const.$ We refer to section 3.1 for an extensive discussion on prior and on critical 'frontier' cases.

5.1 Inferring p_s

If priors are uniform then, Eq. (41) becomes

$$f(p_s\,|\,x,\,n,\,\lambda,p_b) \quad\propto \quad \sum_{n_s,\,n_b} f_{2\mathcal{B}}(x\,|\,n_s,\,p_s\,n_b,\,p_b)\,f(n_b\,|\,\mathcal{P}_{\lambda_b})\,\delta_{n,\,n_s+n_b}\,. \quad (44)$$

Figure 10 gives the result for $x = 9$, $n = 12$, and assuming several hypothesis for λ_b and p_b.

- The upper plot is for $p_b = 0.75$, equal to x/n. The curves are for $\lambda_b = 0$, 1, 2, 4, 6, 8, 10, 12 and 14, with the order indicated (whenever possible) in the figure. If the expected background is null, we recover the simple result we already know. As the expected background increases, $f(p_s)$ gets broader, because the inference is based on a smaller number of objects attributed to the signals and because we are uncertain on the number of events actually due the background. In a very noisy environments ($\lambda_b \approx n$, or even larger), the data provide very little information about p_s and, essentially, the prior pdf (dashed curve) is recovered. Note also that for all values of λ_b the posterior $f(p_s)$ is peaked at $x/n = 0.75$. This is due to the fact that p_b was equal to the observed ratio x/n, therefore, for any hypothesis of n_b attributed to the background, $x_b = p_b\,n_b$ counts are in average 'subtracted' from x (this is properly done in an automatic way in the Bayes formula, followed by marginalization).

- The situation gets more interesting when p_b differs from x/n.

  The middle plot in the figure is for $p_b = 0.25$. Again, the case $\lambda_b = 0$ gives the the pdf we already know. But as soon as some background is hypothesized, the curves start to drift to the right side. That is because high background with low p_b favors large values of p_s.

  The opposite happens if we think that background is characterized by large p_b, as shown in the bottom plot of the figure.

5.2 Inferring n_s and λ_s

The histograms of Fig. 11 show examples of the probability distributions of n_s for $\lambda_b = 4$ and three different hypotheses for p_b. These distributions quantify how much we believe that n_s out of the observed n belong to the signal. [By the way, the number n_b of background objects present in the data can be inferred as complement to n_s, since the two numbers are linearly dependent. It follows that $f(n_b\,|\,x,\,n,\,\lambda_b,\,p_b) = f(n-n_s\,|\,x,\,n,\,\lambda_b,\,p_b)$.]

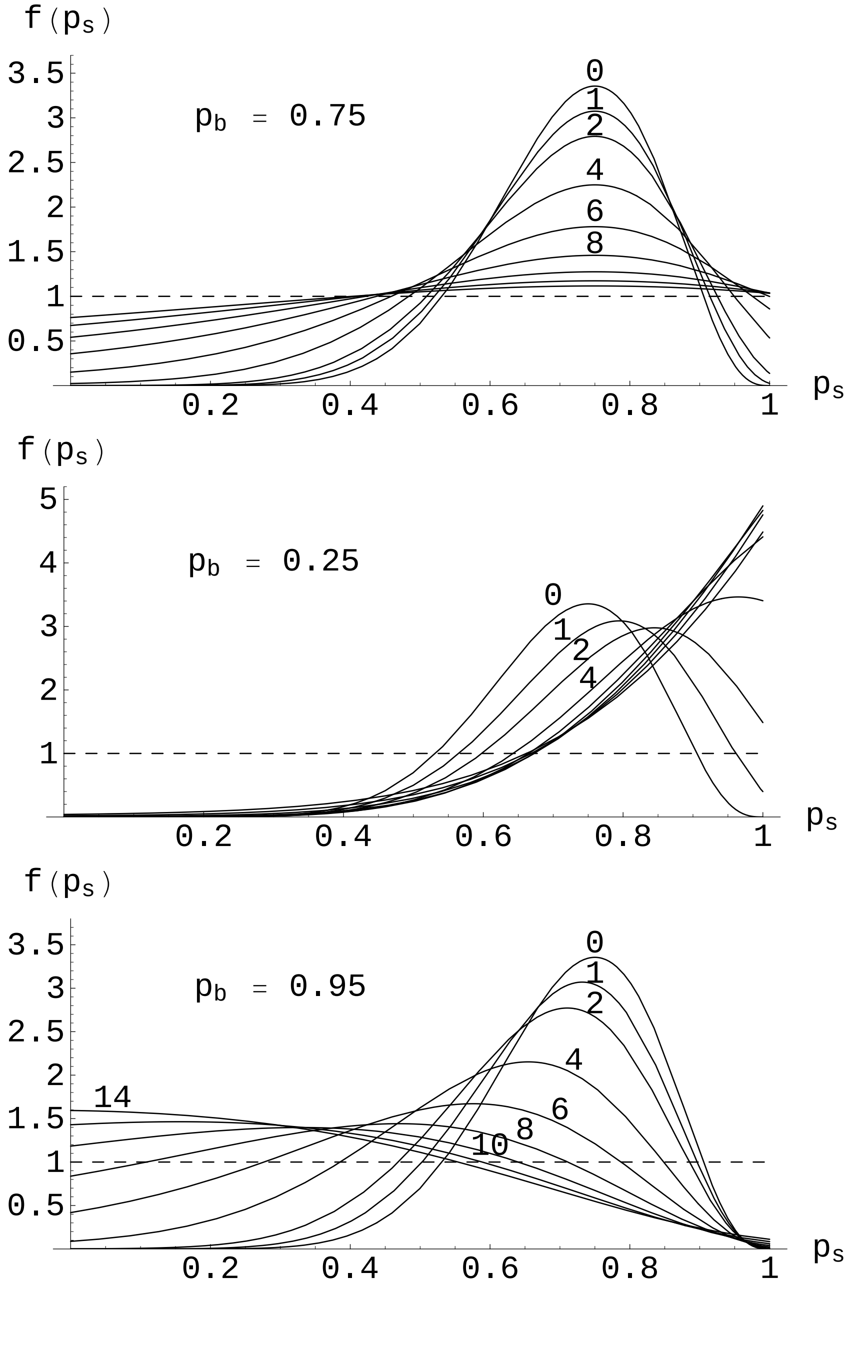


Figure 10: Inference about p_s for $n = 12$ and $x = 9$, depending on the expected background [$\lambda_b = 0,\ 1,\ 2,\ 4,\ 6,\ 8,\ 10,\ 14$, as (possibly) indicated by the number above the lines]. The three plots are obtained by three different hypotheses of p_b.

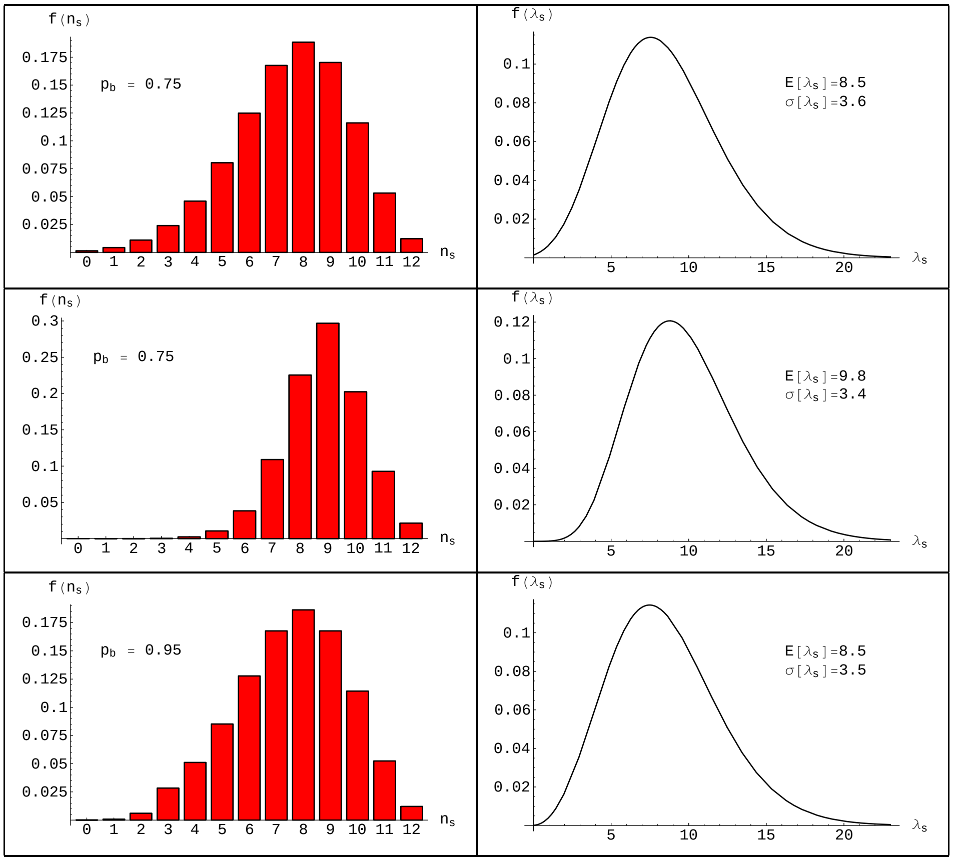


Figure 11: Inference about n_s (histograms) and p_s (continuous lines) for $n = 12$ and $x = 9$, assuming $\lambda_b = 4$ and three values of p_b: 0.75, 0.25 and 0.95 (top down).

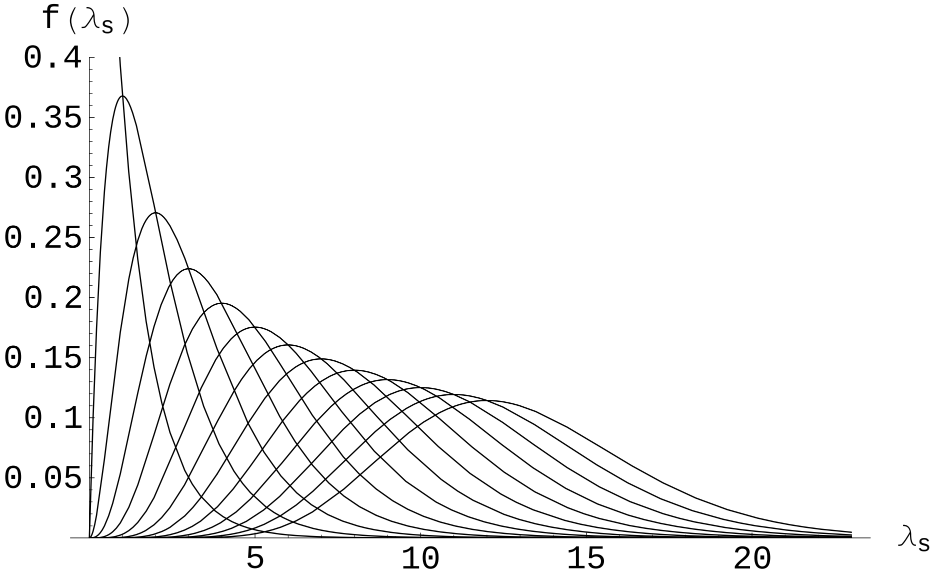


Figure 12: Inference of λ_s depending on the n_s, ranging from 0 to 12 (left to right curves).

A different question is to infer the the Poisson λ_s of the signal. Using once more Bayes theorem we get, under the hypothesis of n_s signal objects:

$$f(\lambda_s \,|\, n_s) \quad \propto \quad f(n_s \,|\, \mathcal{P}_{\lambda_s}) \, .f_0(\lambda_s) \tag{45}$$

Assuming a uniform prior for λ_s we get (see e.g. Ref. [2]):

$$f(\lambda_s \,|\, n_s) \quad = \quad \frac{e^{-\lambda_s}\,\lambda_s^{n_s}}{n_s!}\,, \tag{46}$$

with expected value and variance both equal to $n_s + 1$ and mode equal to n_s (the expected value is shifted on the right side of the mode because the distribution is skewed to the right). Figure 12 shows these pdf's, for n_s ranging from 0 to 12 and assuming a uniform prior for λ_s.

As far the pdf of λ_s that depends on all possible values of n_s, each with is probability, is concerned, we get from probability theory [and remembering that, indeed, $f(\lambda_s \,|\, n_s,\, x,\, n,\, \lambda_b,\, p_b)$ is equal to $f(\lambda_s \,|\, n_s)$, because n_s depends only on λ_s, and then the other way around]:

$$f(\lambda_s \,|\, x,\, n,\, \lambda_b,\, p_b) \quad \propto \quad \sum_{n_s} f(\lambda_s \,|\, n_s)\, f(n_s \,|\, x,\, n,\, \lambda_b,\, p_b)\,, \tag{47}$$

i.e. the pdf of λ_s is the weighted average[2] of the several n_s depending pdf's.

The results for the example we are considering in this section are given in the plots of Fig. 11.

6 Conclusions

The classical inverse problem related to the binomial distribution has been reviewed and extended to the presence of background either only on the number of 'successes', or on the trials themselves. The probabilistic approach followed here allows to treat the problems only using probability rules. The results are always in qualitative agreement with intuition, are consistent with observations and prior knowledge and, never lead to absurdities, like p outside the range 0 and 1.

The role of the priors, that are crucial to allow the probabilistic inversion and very useful to balance in the proper way prior knowledge and evidence from new observations, has been also emphasized, showing when they can be neglected and when they are so critical that it is preferable not to provide probabilistic conclusions.

It is a pleasure to thank Stefano Andreon for several stimulating discussions on the subject.

[2] It follows that all moments of the distribution are weighted averages of the moments of the conditional distribution. Then, expected value and variance of λ_s can be easily obtained from the conditional expected values and variances:

$$\begin{aligned} \mathrm{E}(\lambda_s) &\propto \sum_{n_s} \mathrm{E}(\lambda_s\,|\,n_s)\, f(n_s) \\ \mathrm{Var}(\lambda_s) &\propto \sum_{n_s} [\mathrm{Var}(\lambda_s\,|\,n_s) + \mathrm{E}^2(\lambda_s\,|\,n_s)]\, f(n_s)\,. \end{aligned}$$

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