

## Elementary Sampling theory

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### 3.26, 3.27 The most probable value of $r$

The sequence  $h(r|N, M, n)$  for fixed  $N, M, n$  is unimodal, meaning it first increases, then decreases. To see this we argue as follows.

We want to see whether  $h(r+1|N, M, n)$  is bigger than  $h(r|N, M, n)$ , so we need to compare their fraction to 1. We compute using 3.22

$$h(r+1|N, M, n)/h(r|N, M, n) =$$

$$\frac{(M-r)/(r+1)}{(N-M-n+r+1)/(n-r)} =$$

$$\frac{(r-M)(r-n)}{(r+1)(r+N-M-n+1)} \stackrel{?}{\geq} 1$$

$$(r-M)(r-n) \stackrel{?}{\geq} (r+1)(r+N-M-n+1)$$

$$Mn - (M+n)r \stackrel{?}{\geq} (N-M-n+1) + r(N-M-n+2)$$

$$Mn - (N-M-n+1) \stackrel{?}{\geq} r(N+2)$$

$$\frac{Mn - (N-M-n+1)}{N+2} \stackrel{?}{\geq} r$$

$$\frac{Mn + M + n + 1}{N+2} \stackrel{?}{\geq} r + 1$$

$$\frac{(M+1)(n+1)}{N+2} \stackrel{?}{\geq} r + 1$$

The sequence  $h(r)$  increases while the left hand side is bigger.

Thus denoting by  $r'$  the number  $\frac{(M+1)(n+1)}{N+2}$  we see that if  $r'$  is an integer, then  $h(r)$  increase until  $r = r' - 1$ , then  $h(r' - 1) = h(r')$ , then the  $h(r)$  decrease. If  $r'$  is not an integer, then  $h(r)$  increase until  $h(INT(r'))$ , then decrease.

Remak 1: Note that the expected number of red balls is just the “naive”  $n \frac{M}{N}$  (this is not hard to show using linearity of expectation, see Example 4.2.3 in Blitzstein-Hwang “Introduction to Probability”).

Remark 2: The above result can be restated in the following way: pretend to add one red and one white ball to the urn (for a total of  $N + 2$ ) and draw  $n + 1$  balls from the resulting urn. Compute the “naive” most likely fraction of red balls  $\frac{M+1}{N+2}$  and the “naive” most likely number of red balls  $\frac{(n+1)(M+1)}{N+2}$ . Now subtract 1. This is (up to rounding) the most likely number of red balls drawn in the original procedure. This seems somewhat reminiscent of the correction that putting a beta prior on Bernoulli makes to the posterior expectation, but I have no idea if there is more to this connection than that.

### 3.29 Symmetry of $h(r|N, M, n)$

Combinatorial proof that

$$h(r|N, M, n) = h(r|N, n, M).$$

Remark: This is Theorem 3.4.5 in Blitzstein - Hwang “Introduction to Probability”. See also Theorem 3.9.2.

By definition,  $h(r|N, M, n)$  is computed as follows. Lay down  $N$  balls, labelled  $1, \dots, N$ . Pick the subset  $R_0 = \{1, \dots, M\}$  of them and paint it red. Then pick a subset  $D$  of size  $n$  of all the ball, and compute  $r = |D \cap R_0|$ . The fraction of  $D$ s that give specific answer  $r$  is by definition  $h(r|N, M, n)$ .

Now suppose instead we pick a different subset  $R_1$  of size  $M$  to be red, and repeat the procedure above: pick  $D$ , and compute  $r = |D \cap R_1|$ . We claim that the fraction of  $D$ s that give specific answer  $r$  is still  $h(r|N, M, n)$ . In deed, there exists a permutation of  $\{1, \dots, N\}$  taking  $R_1$  to  $R_0$  (“sort the reds to be first”); the same permutation takes  $D$ s that give  $D \cap R_1 = r$  to those that give  $D \cap R_0 = r$ . Hence there are the same number of  $D$ s in both circumstances.

The above argument means that  $h(r|N, M, n)$  can be also computed as follows. Lay down  $N$  balls, labeled  $1, \dots, N$ . Pick **any** subset  $R$  of them of size  $M$  and paint it red. Then pick a subset  $D$  of size  $n$  of all the ball, and compute  $r = |D \cap R|$ . The fraction of  **$R$ s and  $D$ s** that give specific answer  $r$  is then  $h(r|N, M, n)$ .

But the above procedure remains the same if we exchange  $M$  and  $n$  and rename “paint red” into “pick” and “pick” into “paint red”. Then it computes  $h(r|N, n, M)$ . So the two numbers are equal.

Remark: This view of hypergeometric distribution as giving probabilities of overlap of two subsets (“the red” and “the picked”) removes all time dependence and, in my opinion, sheds a lot of light on the discussion at the end of Section 3.2.

### Exercise 3.2

Modified from stackexchange.

Let  $A_i$  be the statement “all colors except color  $i$  were drawn”. Then

$$P(A_i) = \frac{\binom{N-N_i}{m}}{\binom{N}{m}}$$

This is the number of ways of drawing  $m$  balls from  $N - N_i$  non- $i$  colored balls, divided by the number of ways of drawing  $m$  balls from all  $N$  colored balls. Similarly,

$$P(A_i A_j) = \frac{\binom{N-N_i-N_j}{m}}{\binom{N}{m}}, P(A_i A_j A_k) = \frac{\binom{N-N_i-N_j-N_k}{m}}{\binom{N}{m}} \dots$$

The probability we want is  $1 - P(A_1 + A_2 + A_3 + A_4 + A_5)$ , the probability that no colors will be missing from our draw.

By the sum rule this can be calculated as

$$1 - \sum_i P(A_i) + \sum_{i < j} P(A_i A_j) - \sum_{i < j < k} P(A_i A_j A_k) \dots$$

where the first sum term is over all subsets of 1 color, the second is over all subsets of 2 colors, etc.

This calculation can be done in python like so:

```
def prob_all_colors_drawn(m, N):
    '''
    m is number of balls drawn
    N is a list containing how many of each color is in the urn
    '''
    k = len(N) # number of colors

    total = 1.0 # start with 1
    for i in range(1,k+1):

        # calculate each sum term
        conjunction_prob = 0
        for Ns in combinations(N, i): # for all combinations of i colors
            conjunction_prob += comb(sum(N) - sum(Ns), m, True)/comb(sum(N), m, True)

        # alternately add or subtract the sum term
        total += ((-1)**i)*conjunction_prob
    return total
```

You can modify and run this code on Google Colab, and see a monte carlo approximation of the same problem.

### Exercise 3.3

TODO

I think this depends mainly on your prior distribution  $p(N_1, N_2, \dots, N_k | k)$ . If we can sample from this distribution, and we know your prior over  $k$ , then we can do a monte carlo approximation of  $p(k | \text{colors} = 3) = \sum_{\text{all } N_1 \dots N_k} p(\text{colors} = 3 | k, N_1, N_2, \dots, N_k) p(N_1, N_2, \dots, N_k | k) p(k)$ .

Jaynes seems to think there is only one reasonable prior for this problem, does anyone know what it might be?

### Exercise 3.4

Denote by  $F_i$  be the event “ $i$  is fixed”, and, for any  $I \subset \{1, \dots, n\}$ , denote by  $F_I$  the event “all  $i$  in  $I$  are fixed”, i.e.  $F_I = \prod_{i \in I} F_i$ .

We are looking for  $P(\sum F_i)$ . By inclusion exclusion this is

$$P(\sum F_i) = \sum_{I \subset \{1, \dots, n\}, I \neq \emptyset} (-1)^{|I|+1} P(F_I).$$

For a given subset of size  $k$  probability that it is fixed is  $\frac{(n-k)!}{n!}$ , and there are  $\binom{n}{k}$  such subsets, so the sum over those  $I$  with size  $k$  gives  $(-1)^{k+1} \frac{1}{k!}$ . Plugging this in we obtain

$$h = P(\sum F_i) = \sum_{k=1}^n (-1)^{k+1} \frac{1}{k!},$$

as wanted.

Observe that  $1 - h$  is the value of  $k$ -th order Taylor series for  $e^x$  evaluated at  $x = -1$ , which, as  $k \rightarrow \infty$ , converges to  $e^{-1} = 1/e$ .

### Exercise 3.5

Similarly to 3.4, consider the event  $E_I$  = the bins with labels  $i \in I$  are left empty; then  $P(E_I) = (M - |I|)^N / M^N$  and by inclusion-exclusion  $P(\overline{\sum E_i})$  is

$$\frac{1}{M^N} \sum_{k=0}^M (-1)^k \binom{M}{k} (M - k)^N.$$

Remark: We are computing probability that a function from a set of size  $N$  to a set of size  $M$  is onto. There are  $M^N$  total functions, and the number of surjective ones is  $M!$  times a Stirling number of second kind.

### Some of Exercise 3.6

Remark: If the initial distribution (for  $R_0$ , i.e.  $P(\text{red}) = P(R_0) = p$ ,  $P(\text{white}) = P(\bar{R}_0) = q$ ) were the same as the limit distribution  $\pi$  (formula 3.125,  $\pi(\text{red}) = \lim P(R_k) = \frac{p-\delta}{1-\epsilon-\delta}$ ,  $\pi(\text{white}) = \lim P(\bar{R}_k) = \frac{q-\epsilon}{1-\epsilon-\delta}$ ), this would be a steady state Markov chain, whose time-reverse process is also a Markov chain with transition probabilities  $M_{ij}^r = \frac{\pi_j}{\pi_i} M_{ji}$  (note that for 2 state chains one always has  $M_{ij}^r = M_{ij}$ ). This is precisely condition 3.131. Under this condition it is easy to compute  $P(R_j|R_k)$  with  $j < k$ , and in the 2-state case that we are considering, they would be the same as  $P(R_k|R_j)$  (as in 3.134). However in this exercise the Markov chain starts from the initial distribution that, in general, is not the steady state distribution, so reversing the time produces a process (indexed by negative integers) which is a Markov chain which is not time-homogeneous. Maybe there is still a way to apply general theory of Markov chains to the problem of “backward inference” in this setting; absent that, we proceed by a direct computation (but observe that the reversed process is connected to the limiting behavior of the result, see below).

As usual, all probabilities are conditioned on  $C$ . Equation 3.129 is

$$P(R_k|R_j)P(R_j) = P(R_j|R_k)P(R_k)$$

Equation 3.118 is

$$P(R_k) = \frac{(p-\delta) + (\epsilon+\delta)^{k-1}(p\epsilon - q\delta)}{1-\epsilon-\delta}$$

Finally equation 3.128 is

$$P(R_k|R_j) = \frac{(p-\delta) + (\epsilon+\delta)^{k-j}(q-\epsilon)}{1-\epsilon-\delta}$$

Plugging in

$$P(R_j|R_k) = \frac{(p-\delta) + (\epsilon+\delta)^{j-1}(p\epsilon - q\delta)}{(p-\delta) + (\epsilon+\delta)^{k-1}(p\epsilon - q\delta)} \frac{(p-\delta) + (\epsilon+\delta)^{k-j}(q-\epsilon)}{1-\epsilon-\delta}$$

If both  $j$  and  $k$  go to infinity but  $k-j$  is kept constant, this converges to

$$P(R_{\infty+d}|R_{\infty}) = \frac{(p-\delta) + (\epsilon+\delta)^d(q-\epsilon)}{1-\epsilon-\delta},$$

which is precisely the “reversed process” result (for large  $j$  and  $k$  the influence of the initial distribution not being the stationary one has dissipated; in the 2 state case the reversed process is the same as the forward one, but that’s a

special feature; when the number of states (i.e. colors) is higher limit behavior of “backward inference” is given by the reversed process).