

Second-order Probabilism: Expressive Power and Accuracy

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1 Introduction

As rational agents, we form beliefs about a variety of propositions on the basis of the evidence available to us. But believing a proposition is not an all-or-nothing affair; it is a matter of degrees. We are uncertain, to a greater or lesser extent, about the truth of many propositions since the evidence we possess about them is often fallible. To represent this uncertainty, it is natural to use a probability measure that assigns to each proposition a value between 0 and 1 (also called a degree of belief or credence). This approach—known as *precise probabilism*—models an agent's state of uncertainty (or credal state) with a single probability measure: each proposition is assigned one probability value (a sharp degree of belief). The problem is that a sharp probability measure is not expressive enough to distinguish between intuitively different states of uncertainty rational agents may find themselves in (§2). To avoid this problem, a *set* of probability measures, rather than a single one, can be used to represent the uncertainty of a rational agent. This approach is known as *imprecise probabilism*. It outperforms precise probabilism in some respects, but also runs into problems of its own (§3).

To make progress, this paper argues that the uncertainty of a rational agent is to be represented neither by a single probability measure nor a set of measures. Rather, it is to be represented by a higher-order probability measure, more specifically, a probability distribution over parameter values interpreted as probabilities. Call this view *higher-order probabilism*. We show that higher-order probabilism addresses all the problems and philosophical puzzles that plague both precise and imprecise probabilism (§4 and §5).

Moreover, Bayesian probabilistic programming already provides a fairly reliable implementation framework of this approach.

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2 Precise probabilism

Precise probabilism (PP) holds that a rational agent's uncertainty about a proposition is to be represented as a single, precise probability measure. Bayesian updating regulates how the prior probability measure should change in light of new evidence that the agent learns. The updating can be iterated multiple times for multiple pieces of evidence considered successively. This is an elegant and simple theory with many powerful applications. Unfortunately, representing our uncertainty about a proposition in terms of a single, precise probability measure runs into a number of difficulties.

Precise probabilism fails to capture an important dimension of how our fallible beliefs reflect the evidence we have (or have not) obtained. A couple of stylized examples featuring coin tosses should make the point clear.

No evidence v. fair coin You are about to toss a coin, but have no evidence about its bias. You are completely ignorant. Compare this to the situation in which you know, based on overwhelming evidence, that the coin is fair.

On precise probabilism, both scenarios are represented by assigning a probability of .5 to the outcome *heads*. If you are completely ignorant, the principle of insufficient evidence suggests that you assign .5 to both outcomes. Similarly, if you know for sure the coin is fair, assigning .5 seems the best way to quantify the uncertainty about the outcome. The agent's evidence in the two scenarios is quite different, but precise probabilities fail to capture this difference.

Learning from ignorance You toss a coin with unknown bias. You toss it 10 times and observe *heads* 5 times. Suppose you toss it further and observe 50

heads in 100 tosses.

Since the coin initially had unknown bias, you should presumably assign a probability of .5 to both outcomes if you stick with precise probabilism. After the 10 tosses, you again assess the probability to be .5. You must have learned something, but whatever that is, it is not modeled by precise probabilities. When you toss the coin 100 times and observe 50 heads, you learn something new as well. But your precise probability assessment will again be .5.

These examples suggest that precise probabilism is not appropriately responsive to evidence. Representing an agent's uncertainty by a precise probability measure can fail to track what an agent has learned from new evidence. Precise probabilism assigns the same probability in situations in which one's evidence is quite different: when no evidence is available about a coin's bias; when there is little evidence that the coin is fair (say, after only 10 tosses); and when there is strong evidence that the coin is fair (say, after 100 tosses). In fact, analogous problems also arise for evidence that the coin is not fair. Suppose the rational agent starts with a weak belief that the coin is .6 biased towards heads. They can strengthen that belief by tossing the coin repeatedly and observing, say, 60 heads in 100 tosses. But this improvement in their evidence is not mirrored in the .6 probability they are supposed to assign to *heads*.¹

add reference about sweetening

These problems generalize beyond cases of coin tossing. It is one thing not to know much about whether a proposition is true, for example, whether an individual is guilty of a crime. It is another thing to have strong evidence that favors a hypothesis and equally strong evidence that favors its negation, for example, strong evidence favoring the guilt hypothesis and equally strong evidence favoring the hypothesis of innocence. Despite this difference, precise probabilism would recommend that a probability of .5 be assigned to both hypotheses in either case. Here, too, precise probabilities fail to be appropriately responsive to the evidence.

In addition, evidence can accumulate in a way that does not require changing our initial probability assignments. Suppose that, at first, one's overall evidence favors *A* over *B*. So the probability assigned to *A* should be greater than that assigned to *B*. Next, the agent acquires new evidence. The total quantity of evidence has increased, but suppose this larger body of evidence overall still favors *A* over *B*. So no change in the probabilities seems required. Still, something has changed about the agent's state of uncertainty towards *A* and *B*: the quantity of evidence on which the agent can make their assessment whether *A* is more probable than *B* has become larger. And yet, this change in the quantity of overall evidence is not reflected in the precise probabilities assigned to the propositions *A* and *B*.²

3 Imprecise probabilism

What if we give up the assumption that probability assignments should be precise? Imprecise probabilism (IP) holds that a rational agent's credal stance towards a hypothesis is to be represented by a *set of probability measures*, typically called a *representor* \mathbb{P} , rather than a single measure P . The representor should include all and only those probability measures which

¹Here is another problem for precise probabilism. Imagine a rational agent who does not know the bias of the coin. For precise probabilism, this state of uncertainty should be represented by a .5 probability assignment to the *heads*. Next, the agent learns that the bias towards heads, whatever the bias is, has been slightly increased, say by .001. The addition of this new information is called *sweetening* in the philosophical literature. This sweetening should now make the agent bet on heads: if the probability of *heads* was initially .5, it must be ever so slightly above .5 after sweetening. But, intuitively, the new information should leave the agent equally undecided about betting on heads or tails. After sweetening, the agent still does not know much about the actual bias of the coin.

²Following Keynes, the distinction here is between the *balance* of the evidence—whether the evidence available tips in favor of a proposition or another—and the *weight* of the evidence—the overall quantity of evidence. **ADD REFERENCE TO KEYNES**

are compatible with the evidence (more on this point later).³ It is easy to see that modeling an agent's credal state by sets of probability measures avoids some of the shortcomings of precise probabilism. For instance, if an agent knows that the coin is fair, their credal state would be represented by the singleton set $\{P\}$, where P is a probability measure that assigns .5 to *heads*. If, on the other hand, the agent knows nothing about the coin's bias, their credal state would be represented by the set of all probabilistic measures, since none of them is excluded by the available evidence. Note that the set of probability measures does not represent admissible options that the agent could legitimately pick from. Rather, the agent's credal state is essentially imprecise and should be represented by means of the entire set of probability measures.

So far so good. But, just as precise probabilism fails to be appropriately evidence-responsive in certain scenarios, imprecise probabilism runs in similar difficulties in other scenarios.

Even v. uneven bias: You have two coins and you know, for sure, that the probability of getting heads is .4, if you toss one coin, and .6, if you toss the other coin. But you do not know which is which. You pick one of the two at random and toss it. Contrast this with an uneven case. You have four coins and you know that three of them have bias .4 and one of them has bias .6. You pick a coin at random and plan to toss it. You should be three times more confident that the probability of getting heads is .4. rather than .6.

The first situation can be easily represented by imprecise probabilism. The representor would contain two probability measures, one that assigns .4. and the other that assigns .6 to the hypothesis 'this coin lands heads'. But imprecise probabilism cannot represent the second situation. Since the probability measures in the set are all compatible with the agent's evidence, no probability measure can be assigned a greater (higher-order) probability than any other.⁴

These examples show that imprecise probabilism is not expressive enough to model the scenario of uneven bias. Defenders of imprecise probabilism could concede this point but prefer their account for reasons of simplicity. They could also point out that imprecise probabilism models scenarios that precise probabilism cannot model, for example, a state of complete lack of evidence. In this respect, imprecise probabilism outperforms precise probabilism in expressive power, but also retains theoretical simplicity. Unfortunately, it is questionable whether imprecise probabilism outperforms precise probabilism *all things considered*. In fact, imprecise probabilism suffers from a number of shortcomings that do not affect precise probabilism.

The first shortcoming has not received extensive discussion in the literature, but it is fundamental. Recall that, for imprecise probabilism, an agent's state of uncertainty is represented by those probability measures that are *compatible* with the agent's evidence. The question is, how should the notion of compatibility be understood here? Perhaps we can think of compatibility as the fact that the agent's evidence is consistent with the probability measure in question.

³For the development of imprecise probabilism, see Keynes (1921); Levi (1974); Gärdenfors & Sahlin (1982); Kaplan (1968); Joyce (2005); Fraassen (2006); Sturgeon (2008); Walley (1991). Bradley (2019) is a good source of further references. Imprecise probabilism shares some similarities with what we might call **interval probabilism** (Kyburg, 1961; Kyburg Jr & Teng, 2001). On interval probabilism, precise probabilities are replaced by intervals of probabilities. On imprecise probabilism, instead, precise probabilities are replaced by sets of probabilities. This makes imprecise probabilism more general, since the probabilities of a proposition in the representor set do not have to form a closed interval.

⁴Other scenarios can be constructed in which imprecise probabilism fails to capture distinctive intuitions about evidence and uncertainty; see, for example, (Rinard, 2013). Suppose you know of two urns, GREEN and MYSTERY. You are certain GREEN contains only green marbles, but have no information about MYSTERY. A marble will be drawn at random from each. You should be certain that the marble drawn from GREEN will be green (G), and you should be more confident about this than about the proposition that the marble from MYSTERY will be green (M). In line with how lack of information is to be represented on IP, for each $r \in [0, 1]$ your representor contains a P with $P(M) = r$. But then, it also contains one with $P(M) = 1$. This means that it is not the case that for any probability measure P in your representor, $P(G) > P(M)$, that is, it is not the case that RA is more confident of G than of M . This is highly counter-intuitive.

But mere consistency wouldn't get the agent very far in excluding probability measures, as too many probability measures are strictly speaking still consistent with most observations and data. Admittedly, there will be clear-cut cases: if you see the outcome of a coin toss to be heads, you reject the measure with $P(H) = 0$, and similarly for tails. Another class of cases might arise while randomly drawing objects from a finite set where the true frequencies or objective chances are already known, because the finite set has been inspected. But such clear-cut cases aside, what else? In the end, evidence will often be consistent with almost any probability measure.⁵

A second, well-known problem for imprecise probabilism is belief inertia. Precise probabilism offers an elegant model of learning from evidence: Bayesian updating. Imprecise probabilism, at least *prima facie*, offers an equally elegant model of learning from evidence, richer and more nuanced. It is a natural extension of the classical Bayesian approach that uses precise probabilities. When faced with new evidence E between time t_0 and t_1 , the representor set should be updated point-wise, running the standard Bayesian updating on each probability measure in the representor:

$$\mathbb{P}_{t_1} = \{P_{t_1} | \exists P_{t_0} \in \mathbb{P}_{t_0} \forall H [P_{t_1}(H) = P_{t_0}(H|E)]\}.$$

The hope is that, if we start with a range of probabilities that is not extremely wide, point-wise learning will behave appropriately. For instance, if we start with a prior probability of *heads* equal to .4 or .6, then those measures should be updated to something closer to .5 once we learn that a given coin has already been tossed ten times with the observed number of heads equal 5 (call this evidence E). This would mean that if the initial range of values was $[.4, .6]$ the posterior range of values should be narrower.

Unfortunately, this narrowing of the range of values becomes impossible whenever the starting point is complete lack of knowledge, as imprecise probabilism runs into the problem of belief inertia (Levi, 1980). This problem arises in situations in which no amount of evidence could lead the agent to change their belief state, according to a given modeling strategy. Consider a situation in which you start tossing a coin knowing nothing about its bias. The range of possibilities is $[0, 1]$. After a few tosses, if you observed at least one tail and one heads, you can exclude the measures assigning 0 or 1 to *heads*. But what else have you learned? If you are to update your representor set point-wise, you will end up with the same representor set. For any sequence of outcomes that you can obtain and any probability value in $[0, 1]$, there will exist a probability measure (conditional on the outcomes) that assigns that probability to *heads*. Consequently, the edges of your resulting interval will remain the same. In the end, it is not clear how you are supposed to learn anything if you start from complete ignorance.⁶

Some downplay the problem of belief inertia. After all, if you started with knowing truly nothing, then it is right to conclude that you will never learn anything. Joyce (2010) writes:

⁵Probability measures can be inconsistent with evidential constraints that agents believe to be true. Mathematically, non-trivial evidential constraints are easy to model (Bradley, 2012). They can take the form, for example, of the *evidence of chances* $\{P(X) = x\}$ or $P(X) \in [x, y]$, or *structural constraints* such as “ X and Y are independent” or “ X is more likely than Y .” These constraints are something that an agent can come to accept outright, but only if offered such information by an expert whom the agent completely defers to. But, besides these idealized cases, it is unclear how an agent could come to accept such structural constraints upon observation. There will usually be some degree of uncertainty about the acceptability of these constraints.

⁶Here's another example of inertia, coming from Rinard (2013). Either all the marbles in the urn are green (H_1), or exactly one tenth of the marbles are green (H_2). Suppose your initial credence about these two hypothesis is complete uncertainty with interval. Next, suppose you learn that a marble drawn at random from the urn is green (E). After using this evidence to condition each probability measure in your representor (which initially contains all possible probability measures over the relevant space) on this evidence, you end up with the same spread of values for H_1 that you had before learning E . This holds no matter how many marbles are sampled from the urn and found to be green. This is counterintuitive: if you continue drawing green marbles, even if you started with complete uncertainty, you should become more inclined towards the hypothesis that all marbles are green.

You cannot learn anything in cases of pronounced ignorance simply because a prerequisite for learning is to have prior views about how potential data should alter your beliefs (p. 291)

The upshot is that vacuous priors should not be used and that imprecise probabilism gives the right results when the priors are non-vacuous. Another strategy is to say that, in a state of complete ignorance, a special updating rule should be deployed.⁷

Finally, imprecise probabilism faces a third, deeper problem that does not arise for precise probabilism. As it turns out, it is impossible to define proper scoring rules for measuring the accuracy of a representor. Workable *scoring rules* exist for measuring the accuracy of a single, precise probability measure, such as the Brier score. These rules measure the distance between a rational agent's probability measure and the actual value. A requirement of scoring rules is that they be *proper*: any rational agent will expect their own probability measure to be more accurate than any other. After all, if an agent thought a different probability measure was more accurate, they should switch to it. Proper scoring rules are then used to formulate accuracy-based arguments for precise probabilism. These arguments show (roughly) that, if your precise measure follows the axioms of probability theory, no other non-probabilistic measure is going to be more accurate than yours whatever the facts are. Can the same be done for imprecise probabilism? It cannot. Impossibility theorems demonstrate that no proper scoring rules are available for representor sets (Seidenfeld, Schervish, & Kadane, 2012). So, as many have noted, the prospects for an accuracy-based argument for imprecise probabilism look dim (Campbell-Moore, 2020; Mayo-Wilson & Wheeler, 2016). Moreover, as shown by Schoenfeld (2017), if an accuracy measure satisfies certain plausible formal constraints, it will never strictly recommend an imprecise stance, as for any imprecise stance there will be a precise one with at least the same accuracy.

add reference. Joyce, J. M. (2010). A Defence of Imprecise Credences in Inference and Decision Making, *Philosophical Perspectives* 24, pp. 281–323.

Rafal to develop a response to this point.

4 Higher-order probabilism

Let us take stock. Imprecise probabilism is more expressive than precise probabilism. It can model the difference between a state in which there is no evidence about a proposition (or its negation) and a state in which the evidence for and against a proposition is in equipoise. But imprecise probabilism has its own expressive limitations, for example, it cannot model the case of uneven bias. In addition, imprecise probabilism faces difficulties that do not affect precise probabilism: the notion of compatibility between a probability measure and the evidence is too permissive; belief inertia makes it impossible for a rational agent to learn via Bayesian updating; and no proper scoring rules exist for imprecise probabilism. In this section, we show that higher-order probabilism overcomes the expressive limitations of imprecise probabilism without falling prey to any such difficulties.

Proponents of imprecise probabilism already hinted at the need of relying on higher order-probabilities. For instance, Bradley compares the measures in a representor to committee members, each voting on a particular issue, say the true chance or bias of a coin. As they acquire more evidence, the committee members will often converge on a chance hypothesis.

...the committee members are “bunching up”. Whatever measure you put over the set of probability functions—whatever “second order probability” you use—the “mass” of this measure gets more and more concentrated around the true chance hypothesis. (Bradley, 2012, p. 157)

⁷Elkin (2017) suggests the rule of *credal set replacement* that recommends that upon receiving evidence the agent should drop measures rendered implausible, and add all non-extreme plausible probability measures. This, however, is tricky. One needs a separate account of what makes a distribution plausible from a principled account of why one should use a separate special update rule when starting with complete ignorance.

But such bunching up cannot be modeled by imprecise probabilism alone: a probability distribution over chance hypotheses is needed.⁸ That one should use higher-order probabilities has also been suggested by critics of imprecise probabilism. For example, Carr (2020) argues that sometimes evidence requires uncertainty about what credences to have. Carr, however, does not articulate this suggestion more fully; does not develop it formally; and does not explain how her approach would fare against the difficulties affecting precise and imprecise probabilism. We now set out to do precisely that.

The central idea of higher-order probabilism is this: a rational agent's uncertainty is not single-dimensional and thus cannot be mapped onto a one-dimensional scale like the real line. Uncertainty is best modeled by the shape of a probability distribution over multiple probability measures. Stated more formally, a rational agent's state of uncertainty (or credal stance) towards a proposition X is not represented by a single probability value $P(X)$ between 0 and 1, but by a probability (density) distribution $f_{P(X)}$, where the first-order probability measure $P(X)$ is treated as a random variable. Crucially, this representation is completely general. Unlike the examples used so far, the proposition X is not restricted to chance hypotheses or the bias of a coin. The probability distribution $f_{P(X)}$ assigns a second-order probability to each of the first-order probabilities $P(X)$.

How should these second-order probabilities be understood? It is helpful to think of higher-order probabilism as a generalization of imprecise probabilism. Imprecisers already admit that some probability measures are compatible and others incompatible with the agent's evidence at some point. Compatibility is a coarse notion; it is an all-or-nothing affair. But, as seen earlier, evidence can hardly exclude a probability measure in a definitive manner except in clear-cut cases. Just as it is often a matter of degrees whether evidence supports a proposition, the notion of compatibility between evidence and probability measures can itself be a matter of degrees. On this picture, the evidence justifies different values of first-order probability to various degrees. So, second-order probabilities express the extent to which the first-order probabilities are supported by the evidence.

This higher-order approach at the technical level is by no means novel. Bayesian probabilistic programming languages embrace the well-known idea that parameters can be stacked and depend on each other (Bingham et al., 2021). But, while the technical machinery has been around for a while, it has not been deployed by philosophers to model a rational agent's uncertainty or credal state. Because of its greater expressive power, higher-order probabilism can represent uncertainty in a more fine-grained manner, as illustrated in Figure 1. In particular, the uneven coin scenario in which the two biases of the coin are not equally likely—which imprecise probabilism cannot model—can be easily modeled within high-order probabilism by assigning different probabilities to the two biases.

An agent's uncertainty could—perhaps, should—sometimes be represented by a single probability value. Higher-order probabilism does not prohibit that. For example, there may well be cases in which an agent's uncertainty is aptly represented by the expectation.⁹ But this need not always be the case. If the probability distribution is not sufficiently concentrated around a single value, a one-point summary will fail to do justice to the nuances of the agent's credal state.¹⁰ For example, consider again the scenario in which the agent knows that the bias of the

Rafal to correct technical mistakes in this paragraph (e.g. concept of distribution over probability measures seems incorrect for our purposes)

⁸In a similar vein, Joyce (2005), in a paper defending imprecise probabilism, explicates the notion of weight of evidence using a probability distribution over chance hypotheses. Oddly, representor sets play no central role in Joyce's account of the weight of evidence.

⁹The expectation is usually defined as $\int_0^1 x f(x) dx$. In the context of our approach here, x is the first-order probability of a given proposition, and f is the density representing the agent's uncertainty about x .

¹⁰This approach lines up with common practice in Bayesian statistics, where the primary role of uncertainty representation is assigned to the whole distribution. Summaries such as the mean, mode standard deviation, mean absolute deviation, or highest posterior density intervals are only succinct ways for representing the uncertainty of a given scenario.

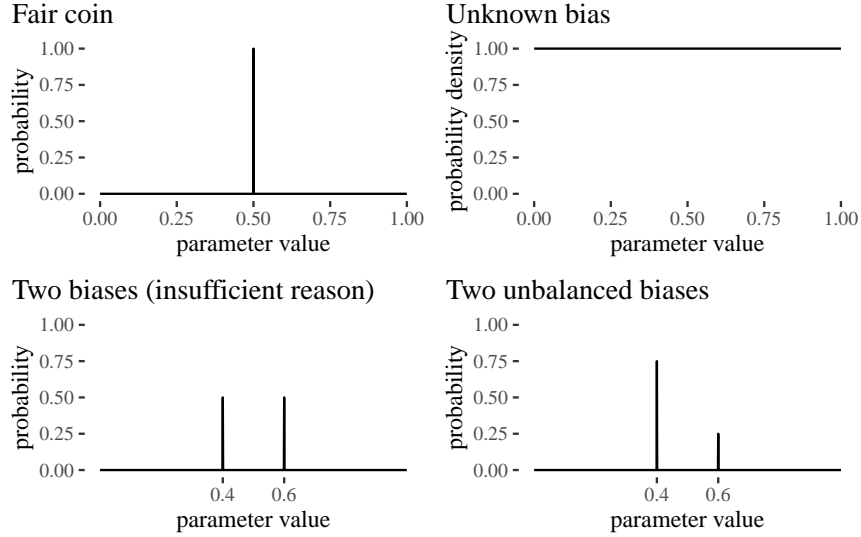


Figure 1: Examples of higher-order distributions for a few scenarios problematic for both precise and imprecise probabilism.

coin is either .4 or .6 but the former is three times more likely. Representing the agent’s credal state with the expectation $P(X) = .75 \times .4 + .25 \times .6 = .45$ would fail to capture the agent’s different epistemic attitudes towards the two biases. The agent believes the two biases have different probabilities, but is also certain the bias is *not* .45.

Besides its greater expressive power in modelling uncertainty, higher-order probabilism does not fall prey to belief inertia or the impossibility of proper scoring rules. Consider a situation in which you have no idea about the bias of a coin. You start with a uniform distribution over $[0, 1]$ as your prior. Observing any non-zero number of heads will exclude 0 and observing any non-zero number of tails will exclude 1 from the basis of the posterior. The posterior distribution will become more centered as the observations come in. This result is a straightforward application of Bayesian updating. Instead of plugging sharp probability values into the formula for Bayes’s theorem, the factors to be multiplied in the theorem will be probability densities (or ratios of densities as needed). Figure 2 illustrates—starting with a uniform prior distribution—how the posterior (beta) distribution changes after successive observations of heads, heads again, and then tails.¹¹

The impossibility of defining proper scoring rules was another weakness of imprecise probabilism. This is a significant shortcoming, especially because proper scores do exist for precise probabilism. Fortunately, one can show that there exist proper scoring rules for higher-order probabilism. These rules can then be used to formulate accuracy-based arguments. In addition, recall the point made by Schoenfield (2017): an accuracy measure will not usually recommend an imprecise stance. This argument fails against imprecise probabilism: there are

¹¹ Assuming independence and constant probability for all the observations, learning is modeled the Bayesian way.

You start with some prior density p over the parameter values. If you start with complete lack of information, p should be uniform. Then, you observe the data D which is the number of successes s in a certain number of observations n . For each particular possible value θ of the parameter, the probability of D conditional on θ follows the binomial distribution. The probability of D is obtained by integration. That is:

$$\begin{aligned} p(\theta|D) &= \frac{p(D|\theta)p(\theta)}{p(D)} \\ &= \frac{\theta^s(1-\theta)^{(n-s)}p(\theta)}{\int (\theta')^s(1-\theta')^{(n-s)}p(\theta') d\theta'}. \end{aligned}$$



Figure 2: As observations of heads, heads and tails come in, extreme parameter values drop out of the picture and the posterior is shaped by the evidence.

cases in which accuracy considerations recommend an imprecise stance (that is, a multi-modal distribution) over a precise one. We will defend these claims in the next section. The argument, however, will be more formal and a bit rough going. The section can be skipped upon first reading without losing track of the main line of the argument.

5 Proper scoring for higher-order probabilism

A scoring rule or inaccuracy score quantifies the distance (inaccuracy) between a probability distribution and the true state of the world. A desirable property that any such score should have is strict propriety. In the precise case, let $I_w(p)$ be an inaccuracy score of a probability distribution p relative to the true state of the world $w \in W$. The score $I_w(p)$ is strictly proper if, for any other probability distribution q different from p , the following holds:

$$\sum_{w \in W} p(w) I_w(p) < \sum_{w \in W} q(w) I_w(p).$$

That is, the expected inaccuracy of p from the perspective of p itself should always be smaller than the expected inaccuracy of p from the perspective of another distribution q . Other common requirements for an inaccuracy score are extensionality (the score should be a function of the probability distribution and the true state) as well as continuity (the score $I_w(p)$ should be a continuous function around p). Well-known results show that the Brier score is an extensional score for precise probabilities that is both proper and continuous.¹²

Can a similar result be established for imprecise probabilities? The answer is likely to be negative, since several hurdles exist. First, it is not straightforward to formulate an intuitively plausible scoring rule for imprecise probabilities. Suppose an imprecise forecast assigns the $[.8, .9]$ interval to the probability of rain, where the true state is ‘rain’. Would the $[.6, .99]$

¹²The Brier score is defined as the squared distance between the true state of the world and the probability forecast, or formally, $(p(w) - V(w))^2$, where $p(w)$ is the probability forecast and $V(w)$ determines if a state $w \in W$ obtains ($V(w) = 1$) or not ($V(w) = 0$). If, for example, the true state of the world is ‘rain’, the forecast ‘rain with .9 probability’ would be more accurate than the forecast ‘rain with .8 probability’. If, on the other hand, the true state were ‘not rain’, the latter forecast would be more accurate.

Is the formal statement of Brier score in the footnote correct? Are probabilities defined over possible worlds or propositions that (partially) describe possible worlds?

interval be more accurate since its .99 upper bound is closer to the true state? Intuitively, the $[\cdot 6, \cdot 99]$ interval should not be more accurate than $[\cdot 8, \cdot 9]$, otherwise the trivial interval $[0, 1]$ would always be more accurate than any other interval. To avoid this outcome, a plausible inaccuracy score for imprecise probabilities could perhaps depend on two things. First, it decreases with the Brier score computed using the side of the interval closer to the true state. Second, the inaccuracy score increases with the size of the interval itself. This blocks the result that the $[0, 1]$ interval is the most accurate. For more details, see Seidenfeld et al. (2012).

Still, however the score for imprecise probabilities is defined, a second hurdle remains. The notion of expected inaccuracy is needed to establish the score's strict propriety, but this notion cannot be easily defined for imprecise probabilities. To see why, let $I_w([p_-, p_+])$ be an inaccuracy score for the interval $[p_-, p_+]$. Its expected inaccuracy from the perspective of the interval $[p_-, p_+]$ itself cannot be easily defined. After all, $I_w([p_-, p_+])$ cannot be multiplied by $[p_-, p_+](w)$, in the way in which $I_w(p)$ can be multiplied by $p(w)$. One option is to evaluate the expected inaccuracy of $[p_-, p_+]$ from the perspective of p_- or p_+ .¹³ But the impreciser would still face a problem. If the expected inaccuracy of the interval $[p_-, p_+]$ is evaluated from the perspective of the precise probabilities p_- or p_+ , finding a proper inaccuracy score that is also continuous turns out to be impossible (Seidenfeld et al., 2012). If, instead, the notion of expected inaccuracy cannot be defined, the very question whether an inaccuracy score for imprecise probabilities can be proper is ill-defined to begin with.

Despite the difficulties that plague imprecise probabilism in defining scoring rules that are proper, here we show that there is a intuitively plausible scoring rule for higher-order probabilities that is both proper and continuous. Building on existing work on this topic (Hersbach (2000), Pettigrew (2012), Gneiting & Raftery (2007)), the higher-order scoring rule we propose is based on well-known measures of divergence between probability distributions. In particular, we rely on the Kullback-Leibler (KL) divergence, which is defined as follows:

$$D_{\text{KL}}(q || p) = \sum_x q(x) \log \left(\frac{q(x)}{p(x)} \right)$$

This is a standard information-theoretic measure of divergence of p from q from the perspective of q . For computational ease, we are using a grid approximation instead of continuous distributions, as in practice we are unable to work with infinite precision.¹⁴ To this end, x will denote the finite vector of discrete states under consideration (by default, we will use equally spaced values between 0 and 1 here).

The goal now is to deploy KL divergence as a measure of inaccuracy of p relative to the true state of the world W , denoted by $I_w^{KL}(p)$. To this end, it is enough for the probability distribution $q(x)$ to play the role of the ‘true’ distribution tracking the true state w ; call it $t(x)$,

$$D_{\text{KL}}(t || p) = \sum_x t(x) \log \left(\frac{t(x)}{p(x)} \right),$$

where to the possible outcomes x , the omniscient distribution t will either assigns 1 (full probability one, if the outcome obtain in w) or 0 (zero probability, if the state does not obtain in w).

That is, it is the KL-divergence of p from the “true” probability distribution which assigns probability 1 to the outcome $V(w)$ and probability 0 to any other outcome,¹⁵ as seen from the perspective of w , the “ground truth” here, conceptually.

¹³So the expected inaccuracy of $I_w([p_-, p_+])$ would equal $\sum_{w \in W} p_-(w) I_w([p_-, p_+])$ or $\sum_{w \in W} p_+(w) I_w([p_-, p_+])$.

¹⁴In the continuous case, we would need to use the so-called differential KL divergence, but our main points wouldn't change otherwise.

¹⁵If you prefer to think in continuous terms, you'd need to use Delta-Dirac distributions here.

Marcello:
I am not
clear what
x is here.
What are
states? Is
the same x
used below
(outcomes)?

The crucial characteristic of a scoring rule that we go after is strict-propiety. Our discussion in this section will be based on an example for the sake of accessibility. We do, however, include the proof of strict propriety of the as an KLD-based inaccuracy measure in the appendix. In a nutshell, the argument demonstrates that for a second-order discretized probability mass p over a parameter space $[0, 1]$, with the actual probability denoted as θ , the Kullback-Leibler divergence of p from the indicator distribution of θ (which assigns 1 to θ and 0 to all other parameter values in the parameter space) is expressed as $\mathcal{J}_{D_{\text{KL}}}^2$.¹⁶ This serves as a demonstration of the strict propriety of the inaccuracy measure: each p anticipates itself to be the least inaccurate distribution.¹⁷

Let's now compare, and show on an example the difference between our approach, which uses KL-divergence and a fairly standard approach represented by CRPS¹⁸. todo{Say that it is used for example by Konek}. We will start by providing a definition of another inaccuracy measure, Cramer-Von-Mises. In the discretized version, it is defined as:¹⁹

$$D_{\text{CM}}(p, q) = \sum_x |P(x) - Q(x)|^2$$

where P is the cumulative probability corresponding to the probability distribution p . This inaccuracy measure can be used to define the inaccuracy of p with respect to w as follows:

$$I_{\text{CRPS}}(p, w) = \sum |P(x) - \mathbf{1}(x \geq V(w))|^2$$

where:

$$\mathbf{1}(x \geq V(w)) = \begin{cases} 1 & \text{if } x \geq V(w) \\ 0 & \text{o/w.} \end{cases}$$

Cramer-Von-Mises can also be used to define the expected inaccuracy, analogously to the case of KL-divergence.

At this point, a friend of point estimates might insist, *but who do we have to pay attention to all these possible values, we can just run with point estimates!* To illustrate why the move to consider values in-between without prior integration, we will go over an example in which there are only two possible low-level outcomes, tails and heads, *and we take the distribution's expected values as the probabilities they assign to those outcomes. On this approach, instead of calculating inaccuracy w.r.t. to all possible hypotheses about the coin's bias, we first average out, and proceed in terms of probabilities of the two outcomes only instead. That is, we use:

¹⁶The argument generalizes to parameter spaces that correspond to probabilities of multiple propositions which are Cartesian products of parameter spaces explicitly used in the argument in this section.

¹⁷The argument has four key moves. Consider a given "point probability", i.e. the unique x such that $V(w) = x$. Call this parameter θ .

1. we show that the inaccuracy of p w.r.t. θ can be reduced to $-\log_2 p(\theta)$,
2. then we show that the expected inaccuracy of p from the perspective of p is the entropy of p , $H(p)$,
3. next we show that the inaccuracy of q from the perspective of p is the cross-entropy $H(p, q)$,
4. we then plug in the well-established result that cross-entropy is strictly larger than entropy as soon as $p \neq q$.

¹⁸Note that there are no readily computable solutions to the integral used in the definition of CRPS, although it can sometimes be evaluated in closed form (Gneiting & Raftery, 2007, p. 366).

¹⁹In the continuous case, this measure is defined as the area under the squared Euclidean distances between the corresponding cumulative density functions That is, $D_{\text{CM}}(p, q) = \int_0^1 |P(x) - Q(x)|^2 dx$. Looking at cumulative densities ensures that all densities are considered on the same scale.

$$\mathbb{E}_{\text{binary}}(p, q) = I(p, \text{heads})\mathbb{E}q(\text{heads}) + I(p, \text{tails})\mathbb{E}q(\text{tails}).$$

We will start by reflecting on this binary approach, but we will soon see that it runs into difficulties. To fix ideas, consider a variation of a scenario by Schoenfield (2017). A rational agent is invited to engage in a bet by an opponent who has a representative bag of coins coming from a factory where the distribution of bias among the coins produced, the true generative process, is known. It is a mixture of two normal distributions centered at .3 and at .5, both with standard deviation of .05. The opponent randomly selects one of these coins and flips it. The rational agent knows all the details of this set-up.

What credal state should the rational agent form in response? Consider three out of many possible options: first, a *faithful bimodal* distribution centered at .3 and .5; second, a *unimodal* distribution centered at .4; third, a *wide bimodal* distribution centered at .2 and .6. The three options are depicted in Figure 3. All of them have expected values at .4 rounded to four digits.

Denote these distributions as b , c , and w correspondingly. Now consider what happens if we think of expected inaccuracy of these distributions assuming there are only two possible true outcomes, conceptualized as either heads (H) or tails (T). Whatever our inaccuracy measure, we will have six inaccuracy scores of the form $I(\text{distribution}, \text{outcome})$, where outcome is one of two omniscient distributions that give all weight to either heads or tails. In the calculation of expected values, these will need to be multiplied by the probability of H and the probability of T . However, none of these distributions assigns such a probability. They all assign probabilities to different values of coin bias, but not to different *outcomes* directly. Suppose we do what seems to be the natural way to go for the friends of point estimation and obtain such probabilities by taking expected values of the form $\mathbb{E}_{\text{distribution}}(H) = \sum(x * \text{distribution}(x))$, where x are the values on the discretized grid (as in our example these expectations are pretty much the same, we can simply take $\text{distribution}(H)$ to be .4):

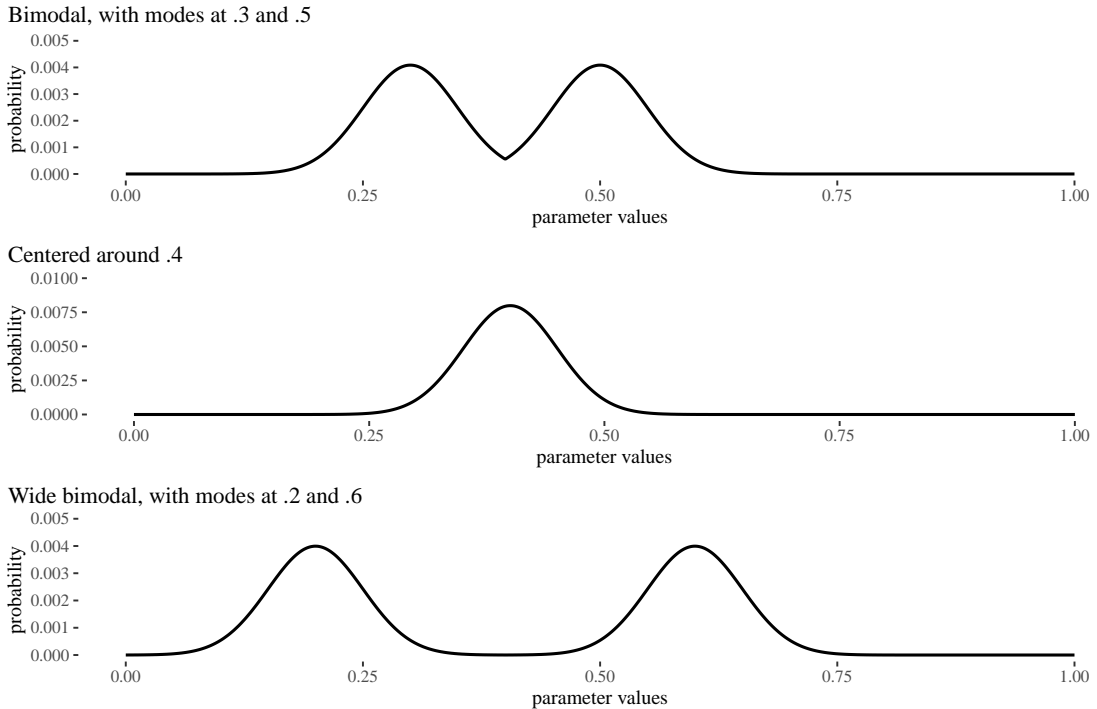


Figure 3: Three distributions in a vague EMS scenario. The distributions are built from normal distributions with standard deviation .05, the bimodal ones are joint in the middle. All of them have expected values $\approx .4$.

Table 1 displays the accuracy scores assigned by CRPS assuming the true outcome is *heads* (CRPS1) and assuming the true outcome is *tails* (CRPS0), and analogously for the KL divergence. Then, we obtain the expected inaccuracies by averaging the two scores by using the point probability of heads set to .4. These result in ExpCRPS and ExpKL respectively, note that since the probability of heads is the same on all the distributions, those are expected values from the perspective of each of the measures; changing the perspective in this example doesn't change the expected inaccuracy.

distribution	CRPS1	CRPS0	KLD1	KLD0	ExpCRPS	ExpKLD
bimodal	534.7305	334.9305	80.06971	33.90347	414.8505	52.36997
centered	571.2192	371.4192	110.84220	53.13440	451.3392	76.21752
wide bimodal	485.4052	285.6177	54.13433	19.50965	365.5340	33.35974

Table 1: CPRS and KLD inaccuracies of the three distributions to the TRUE and FALSE omniscient functions, with expected inaccuracies, calculated using the shared point estimates of the probabilities of heads and tails.

As it turns out, the expected CRPS score recommends the wide bimodal distribution as the most accurate (or least inaccurate). The KL divergence from the omniscient measure makes the same recommendation.²⁰ This is counterintuitive, because the faithful bimodal seems the one most appropriately evidence-responsive. The unimodal distribution, while centering on the expected value, gets the chances wrong, and the wide bimodal has its guesses too close to the truth values and too far from the known chances. But there is a further problem. While the wide bimodal distribution expects itself to be the least inaccurate, the other distributions also expect the wide bimodal to be the least inaccurate. This indicates that in this setting the CRPS score and the KL divergence are not proper scores, as they allow for cases of some distributions recommending other distributions as less accurate, whatever the true state of the world turns out to be.

What are we to make of this result? Note that the three distributions share the same expected value .4. The latter is then used in the calculations of the expected inaccuracies for both CRPS and KL with the assumption there are two possible outcomes with respect to which expected inaccuracy calculations are to be made. This approach, however, runs against the spirit of our enterprise. If expected values are often not good representations of a rational agent's uncertainty, it should not be surprising that relying on them fails to deliver plausible expected accuracy scores. By reducing each of the distributions' stance towards heads to a single point value .4, we've effectively washed out key information. So the question is, how can we adequately account for the complexity of a rational agent's credal state in formulating a proper accuracy score?

Our proposal is that rather than measuring inaccuracy in relation to "true states of the world" conceptualized as two omniscient credences that peak at either 0 or 1 and averaging using expected values of the distributions, we should instead utilize a set of n potential true probability hypotheses (ideally, going continuous, but we're working with a discrete grid of $n = 1000$ possible coin biases in this paper). We then compute all the inaccuracies with respect to each of these n values represented by "omniscient" distributions (or true chance hypotheses) and determine the expected inaccuracy scores using the entire distributions rather than relying solely on the expected values of the distributions.

For the three distributions under consideration, the accuracy scores calculated using CRPS and KL divergence with respect to omniscient distributions corresponding to various values of

²⁰This indicates that the choice of the evaluation metric is not the cause of the recommendation.

x are given by Figure 4. The expected inaccuracies of the distributions from their perspective are given by Table 2. The results now match our common sense: each distribution recommends itself. So, once we pay attention to the whole range of possibilities, the CRPS score and KL divergence are now proper scores.

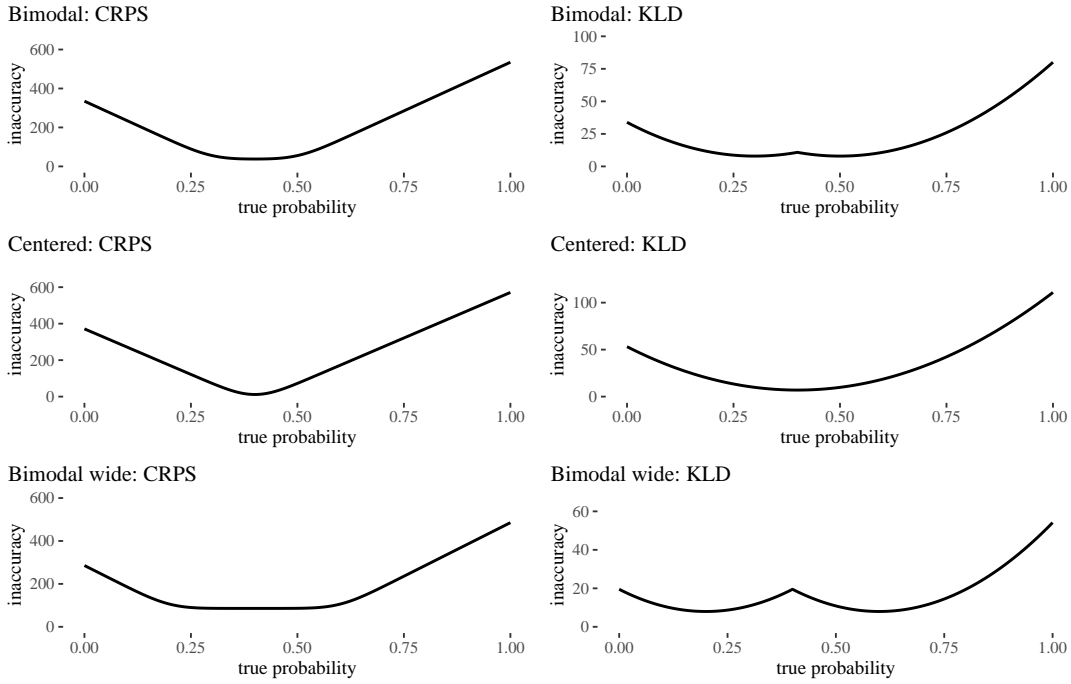


Figure 4: CLPSR and KL divergence based inaccuracies vs (omniscient functions corresponding to) n true probability hypotheses for the three distributions discussed in this section.

	CPRS			KLD		
	bimodal	centered	wide bimodal	bimodal	centered	wide bimodal
bimodal	64.670	78.145	88.380	8.577	10.655	11.336
centered	41.657	28.181	85.911	9.239	7.690	15.627
wide bimodal	137.699	171.719	113.989	11.541	19.231	8.689

Table 2: Expected inaccuracies of the three distributions from their own perspectives. Each row corresponds to a perspective.

One important difference transpires between using CRPS and KLD. Notice how for chance hypotheses between the actual peaks the inaccuracy remains flat. This seems to be an artifact of choosing a squared distance metric. If instead we go with a more principled, information-theory-inspired KL divergence, inaccuracy in fact jumps a bit for values in between the peaks for the bimodal distributions, which seems intuitive and desirable. This seems to be a reason to prefer a KL-based accuracy measure.

One question remains: how does the framework capture the idea that it is the bimodal distribution that seems more adequate than the others? One way to interpret that is by looking at inaccuracy concerning chance hypotheses given by the testimonial evidence. In this case, these are H_3 , where the true chance is 0.3, and H_5 , where the true chance is 0.5. You can find the specific inaccuracies for them in Table 3. To make sure that this favorable outcome isn't

due to not using pointed credences, we can redo the calculations using the pointed version. In the pointed version, all the focus is on 0.4, or the weight is evenly divided between 0.3 and 0.5, or between 0.2 and 0.6. As anticipated, when we consider inaccuracy, both of these setups recommend the bimodal version (Table 4).

	CRPS		KLD	
	H3	H5	H3	H5
bimodal	55.475	55.378	7.935	7.935
centered	72.281	72.090	9.836	9.825
wide bimodal	86.230	86.223	10.871	10.882

Table 3: CRPS and KLD inaccuracies of the three distributions with respect to the two hypotheses. On both inaccuracy measures the bimodal distribution dominates the other two.

	CRPS		KLD	
	H3	H5	H3	H5
pointed bimodal	49.75	49.75	1.00	1.00
pointed centered	100.00	100.00	16.61	16.61
pointed wide bimodal	99.75	99.75	16.61	16.61

Table 4: CRPS and KLD inaccuracies of the three-pointed distributions with respect to the two hypotheses.

6 Evidence aggregation: the simple case

Rational agents are often tasked with aggregating pieces of evidence and assessing their value relative to a hypothesis. In this and the next section, we examine the question of how multiple items of evidence should be evaluated together. This question raises novel difficulties for both precise and imprecise probabilism. We show how higher-order probabilism can handle them.

For the precise probabilist, a natural measure of the value of the evidence is the likelihood ratio. This ratio is relative to a pair of competing hypotheses, say H and its negation $\neg H$ (though the two hypotheses need not be one the negation of the other). Relative to these hypotheses, the likelihood ratio of a single piece of evidence E is the probability of E given H divided by the probability of E given $\neg H$, or in short, $\frac{P(E|H)}{P(E|\neg H)}$. Degrees of evidential value (or support, strength) can be expressed as follows:

the higher $\frac{P(E|H)}{P(E|\neg H)}$ (if greater than one), the more strongly E supports H .

The value of the evidence increases whenever $P(E|H)$ increases or whenever $P(E|\neg H)$ decreases. The higher $P(E|H)$, the better the evidence at tracking H (a true positive); the lower $P(E|\neg H)$, the better the evidence at avoiding $\neg H$ (a true negative). If the probability of E is the same given hypothesis H as given its negation, that is, the likelihood ratio equals one, the evidence would have no value for H .

Likelihood ratios can also be used for assessing the value of multiple pieces of evidence in the aggregate, again relative to a pair of hypotheses of interest. In the simplest case (for more complex cases, see the next section), multiple items of evidence all bear on the same hypothesis. Then, to obtain their combined evidential value, it is enough to multiply their

individual likelihood ratios.

$$\frac{P(E_1 \wedge E_2 \dots E_k | H)}{P(E_1 \wedge E_2 \dots E_k | \neg H)} = \frac{P(E_1 | H)}{P(E_1 | \neg H)} \times \frac{P(E_2 | H)}{P(E_2 | \neg H)} \times \dots \times \frac{P(E_k | H)}{P(E_k | \neg H)}$$

The equality holds provided E_1, E_2, \dots, E_k are probabilistically independent conditional on hypothesis H and its negation. Think, for example, at several diagnostic tests performed by independent laboratories or independent witnesses in a trial testifying about the same issue.

To see how likelihood ratios can be deployed, it is worth working through a specific case. In a murder case, the police recover two items of trace evidence, both against the defendant. First, hair found at the crime scene matches the defendant's hair; call this evidence 'hair.' Second, the fur of the defendant's dog matches the fur found in a carpet wrapped around one of the bodies; call this evidence 'fur.'²¹ The two matches favor the hypothesis that the defendant (and the defendant's dog) must be the source of the crime traces; call this hypothesis 'source'. If the two matches are independent lines of evidence (conditional on the source hypothesis and its negation), their likelihood ratios can be multiplied:

$$\frac{P(\text{fur} \wedge \text{hair} | \text{source})}{P(\text{fur} \wedge \text{hair} | \neg \text{source})} = \frac{P(\text{fur} | \text{source})}{P(\text{fur} | \neg \text{source})} \times \frac{P(\text{hair} | \text{source})}{P(\text{hair} | \neg \text{source})}$$

So far so good. But how do we fill in the precise probabilities? The numerators can be equated to one: if the defendant is a contributor, the laboratory will declare a match for sure. This is a simplification, but it will do for our purposes. To fill in the denominators, a trial expert will provide so-called match probabilities. They express the likelihood that, by coincidence, a random person (or a random dog) who is not a contributor would still match. The match probabilities are approximated by counting how many matches are found in a representative sample of the human population (or the canine population). Suppose the matching hair type occurs 0.0253 times in a reference database, and the matching dog fur type occurs 0.0256 times in a reference database (more on how these numbers are calculated soon). These frequencies can fill in the match probabilities. Putting everything together:

$$\frac{P(\text{dog} | \text{source})}{P(\text{dog} | \neg \text{source})} \times \frac{P(\text{hair} | \text{source})}{P(\text{hair} | \neg \text{source})} = \frac{1}{0.0252613} \times \frac{1}{0.025641} = 1543.862069$$

The resulting ratio is large. The two matches, combined, strongly favor the source hypothesis.

This is the story about evidence aggregation told by the precise probabilist. But this story misses something crucial. As it happens, the match probability for hair evidence is based on 29 matches found in a sample database of size 1148, while the match probability for the dog evidence is based on finding two matches in a smaller database of size 78. The relative frequencies are about .025 in both cases, but the two samples differ in size. The smaller the sample, the greater the uncertainty about the match probabilities. So, for individual pieces of evidence, simply reporting the exact numbers makes it seem as though the evidential value of the matches is the same, but actually it is not.²² In the aggregate, multiplying the individual likelihood ratios further washes away this difference.

A better alternative is easily available: the evaluation of multiple items of evidence should take into account higher-order uncertainty. Figure 6 (upper part) depicts higher-order probability distributions of different match probabilities given the sample data—the actual number

²¹The hair evidence and the dog fur evidence are stylized after two items of evidence in the notorious 1981 Wayne Williams case (Deadman, 1984b, 1984a).

²²The match probabilities in the Wayne Williams case on which our running example is based were 1/100 for the dog fur, and 29/1148 for Wayne Williams' hair. Match probabilities have been slightly but not unrealistically modified to be closer to each other in order to make a conceptual point: the same first-order probabilities, even when they sound precise, may come with different degrees of second-order uncertainty.

of matches found in the sample databases. As expected, some random match probabilities are more likely than others, and since the sizes of the two databases are different, the distributions have different spreads: the smaller the database the greater the spread, the greater the uncertainty about the match probability. In light of this, Figure 6 (lower part) depicts the probability distribution for the joint match probability associated with both items of match evidence, hair and fur evidence. The aggregate value of the two pieces of match evidence, then, is given by a distribution over possible likelihood ratios. The shape of this distribution conveys the degree of higher-order uncertainty about the value of the aggregate evidence. **Marcello (Rafal/Nikodem to add): Can we have a formula of how the two matches are combined in the higher-order approach? In precise probabilism, you multiply the individual LRs, in higher-order probabilism, what do we do formally? Can we also have a distribution of likelihood ratios? What happens if both numerator and denominator in the LR are distributions?**

add distributions of LRs figure

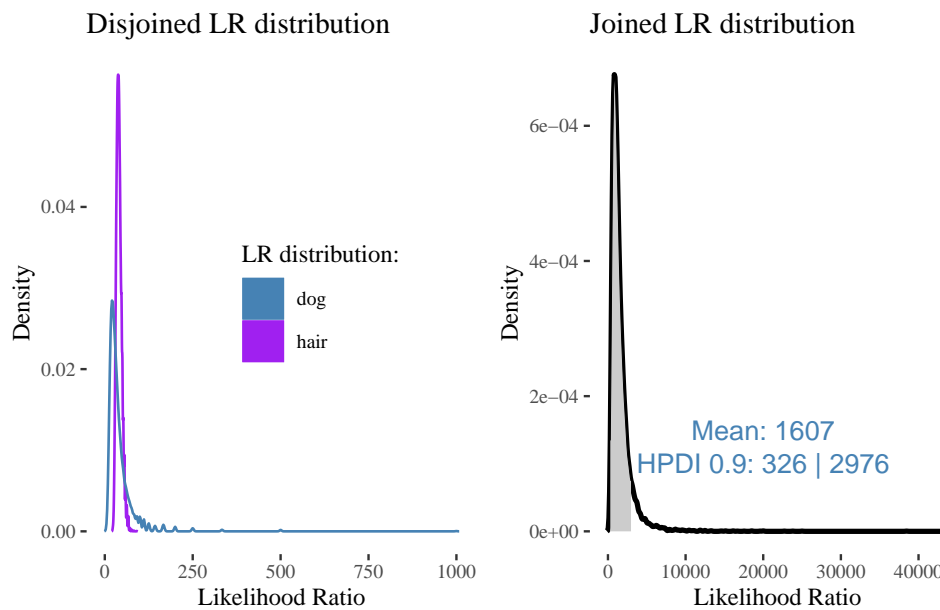


Figure 5: Distributions of dog and hair likelihood ratios and the resulting joint likelihood ratio. Created with the samples from beta distributions. Shaded area on the second one represents HPDI with 0.9 credibility.

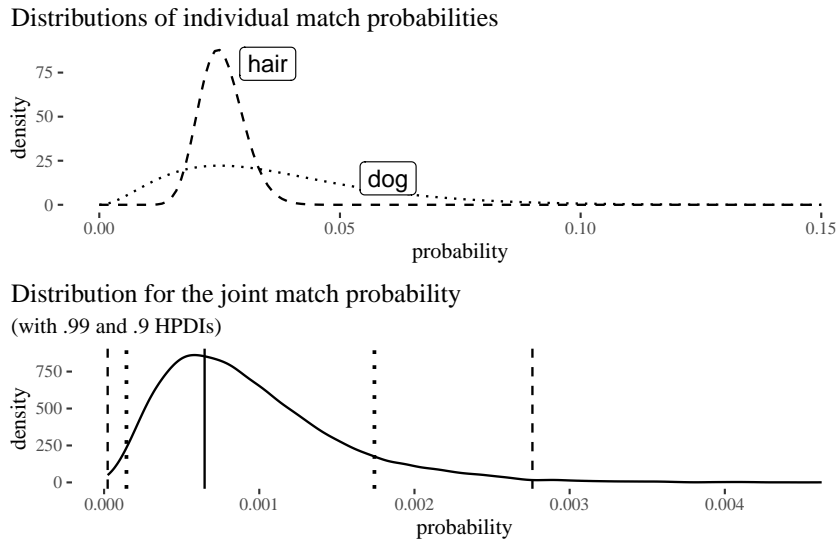


Figure 6: Beta densities for individual items of evidence and the resulting joint density with .99 and .9 highest posterior density intervals, assuming the sample sizes as discussed and independence, with uniform priors.

The precise probabilist might insist that the value of the evidence—one item or multiple items of evidence—is naturally captured by a precise likelihood ratio. In our running example, this ratio equals one over the first-order match probability, and our best assessment of this first-order probability is still the relative frequency of matches found in the database, whether large or small. If this is right, our best assessment of the match probabilities for both fur and hair evidence should be about .025, based on the relative frequencies 2/78 and 29/1148. If we were to bet whether a dog or a human picked at random would have the matching fur or hair type, our odds should be .025 no matter the size of the database. This argument has some bite when evaluating single items of evidence. In fact, the expected values of the match probabilities for hair and match evidence—based on the higher-order distributions in Figure 6 (upper part)—still end up being about .025. If, as the precise probabilist assumes, first-order probabilities is all we should care about, going higher-order would seem a needless complication.

This line of reasoning, however, breaks down when evaluating two or more items of evidence. What should our betting odds be for the proposition that a human and a dog, both picked at random, would have the matching fur and hair type in question? For the precise probabilist, the answer is straightforward: on the assumption of independence, it is enough to multiply the .025 individual match probabilities and obtain a joint match probability of 6.4772626×10^{-4} . The higher-order probabilist will proceed differently. In assessing first-order match probabilities, they will retain information about higher-order uncertainty as much as possible. This can be done in two steps: first, aggregate the higher-order distributions for the two match probabilities and obtain a higher-order probability distribution for the joint match probability (see Figure 6); next, to obtain our best assessment of the first-order joint match probability, take the expected value of this latter distribution. The higher-order probabilist will assign 9.381365×10^{-4} to the joint match probability, a value greater than what the precise probabilist would assign.

So, the higher-order and precise probabilist will disagree about the betting odds for the proposition that a human and a dog, both picked at random, would have the matching fur and hair type. The disagreement will become even starker as a larger number of independent items of evidence are evaluated.²³ Who should be trusted? Since the higher-order probabilist

²³Consider the simple case of independent items of evidence whose individual match probabilities are .025. For three, five and seven items of evidence, the joint match probabilities would be: 1.25×10^{-4} , $3.125 \times$

takes into account more information—that is, the higher-distributions—there is good reason to think that the higher-order probabilist should be trusted more than the precise probabilist.²⁴

Imprecise probabilism will also run into its own problems when assessing the value of aggregate evidence. Recall that the probability measures in the representor set are those compatible with the evidence. The problem is that almost any random match probability will be compatible with any sample data—with any number of matches found in a reference database. This point should be familiar from the earlier discussion. Think by analogy to coin tossing: even a coin that has a .99 bias toward tails could come up heads on every toss. This series of outcomes is unlikely, but possible. Similarly, a hair type that has a match probability extremely small could still be found several times in a sample population. So, it is not clear how to proceed if one takes seriously the binary notion of compatibility. Imprecise probabilism is too permissive because almost any match probability will be compatible with the data.

Another option for the imprecise probabilist is to rely on reasonable ranges of match probabilities. Suppose these ranges are (.015, .037) (.002, .103), for hair and fur evidence respectively in our original case.²⁵ As expected, the range is wider for dog fur match evidence than hair match evidence: the uncertainty about the dog fur match probability is greater since the sample database was smaller. This is a desirable feature of the interval approach. Now, to assess the joint uncertainty, it is enough to focus on what happens at the edges of the two intervals. Reasoning with representor members at the edges of the intervals will yield the most extreme probability measure the impreciser is committed to, the worst-case and best-case scenarios. We end up with a new range for the joint match probabilities, (.00003, .003811).²⁶ The corresponding likelihood ratios could be as high as 3.333333×10^4 or as low as 262.3983. The two matches could be much stronger or much weaker evidence than previously thought.

Using plausible ranges for the match probabilities leaves the impression that any value in the interval is just as good as any other. Perhaps we should pick the middle value as representative

10^{-7} and 7.8125×10^{-10} (for the precise probabilist) and 5.3363868×10^{-4} , 1.6754185×10^{-5} and 9.9986742×10^{-7} (for the higher-order probabilist, based on small databases of size 20).

²⁴As a further illustration of this point, consider a couple of variations of our running example. First, suppose the match probabilities associated with two matches are both set to .05, since they are based on the following relative frequencies: one match occurs in a dog fur database and one match occurs in a human hair database, where both databases are small, say of size 20. By multiplying the individual $1/.05$ likelihood ratios associated with the two matches, their evidential value against the defendant would seem quite strong: $1/.05 \times 1/.05 = 400$. But the match probabilities are based on frequencies resulting from small databases, so their evidential value should be rather weak. Precise probability here seems to exaggerate the aggregate value of the evidence. Following higher-order probabilism, the joint likelihood ratio would be 237.8675988, a significantly smaller value. On the other hand, if the same .05 match probabilities were based on larger databases, the evidential value of the two matches should be correspondingly greater, but precise probabilism would make no difference. If, for example, 1,000 hair and fur matches are found in databases of size 20,000, the higher-order probabilist would assign 441.2059925 to the joint likelihood ratio, a much greater value than before. This outcome agrees with our intuitions.

²⁵These are 99% credible intervals starting with uniform priors. A 99% credible interval is the narrowest interval to which the expert thinks the true parameter belongs with probability .99. For a discussion of what credible intervals are, how they differ from confidence intervals, and why confidence intervals should not be used, see Kruschke (2015).

²⁶Redoing the calculations using the upper bounds of the two intervals, .037 and .103, yields the following:

$$P(\text{dog} \wedge \text{hair} | \neg \text{source}) = .037 \times .103 = .003811.$$

This number is around 5.88 times greater than the original estimate. Given this number, the two matches are much weaker evidence for the source hypothesis than previously thought. The calculation for the lower bounds, .015 and .002, yields the following:

$$P(\text{dog} \wedge \text{hair} | \neg \text{source}) = .015 \times .002 = .00003$$

This number is around 0.46 times lower than the original estimate. Given this number, the two matches are much stronger evidence for the source hypothesis than previously thought.

of the interval. However, relying on the entire interval or the middle value will misrepresent the evidence. To see why, consider again Figure 6 (lower part) which depicts the probability distribution for the joint match probability. Interestingly, this distribution is not symmetric. So the most likely value (and the bulk of the distribution, really) does not lie in the middle between the edges. Therefore, only relying on the edges—or taking central values as representative of the interval—can lead to overestimating or underestimating the probabilities at play.²⁷

Another problem in taking intervals as representative of the value of the evidence is that they will tend to widen as more items of evidence are evaluated. The size of the likelihood ratio interval was initially 39.64 (hair evidence) and 490.29 (fur evidence). After aggregating the two items of evidence, the likelihood ratio interval widened to 3.3070935×10^4 . The size of the match probability interval was initially -0.022 (hair evidence) and -0.101 (fur evidence). After aggregating the two items of evidence, the match probability interval narrowed to -0.003781. Posterior interval (starting with 1:1 prior odds) was initially 0.0209015 (hair evidence) and 0.0913857 (fur evidence). After aggregating the two items of evidence, the posterior interval narrowed to 0.0037665.

All in all, precise and imprecise probabilism do not fare well in modelling the value of evidence in the aggregate. Instead, the evaluation of multiple items of evidence should take into account higher-order uncertainty (as illustrated in Figure 6). Whenever probability distributions for the probabilities of interest are available (and they should be available for match evidence and many forms of scientific evidence whose reliability has been studied), those distributions should be reported for assessing the value of the evidence. This approach avoids hiding actual aleatory uncertainties under the carpet. It also allows for a more balanced assessment of the evidence, whereas using point values or intervals may exaggerate or underestimate the value of the evidence.

A couple of clarifications are in order. First, the problem we are highlighting is not confined to match evidence. Say an eyewitness testifies against the defendant: they saw the defendant near the crime scene at the relevant time. To assess the value of this testimony, one should know something analogous to the match probability: if the defendant was not there, how probable is it that the witness would still say the defendant was there? Or suppose a medical test for a disease turns out positive. Here again, to assess the evidential value of the positive test, one should know how probable it is that the test would still turn out positive even when a patient is actually negative. And so on. These false positive probabilities are usually derived from sample-based frequencies in surveys or experiments: how often witness misidentify people; how often tests misdiagnose; etc. So, depending on the sample size, the false positive probabilities will have different degrees of uncertainty, and the latter should be taken into account when evaluating eyewitness testimonies, diagnostic test results, and many other forms of evidence. At the same time—and this is the second clarification—this discussion is not meant to suggest that the problem we are highlighting is confined to differences in sample size; it is broader than that. Probabilities can be subject to uncertainty for other reasons, for example, when they are derived from a probability model for which there is little support, or when the sample size is large but unrepresentative. So, in short, the problem of higher-order uncertainty is widespread and goes beyond match evidence and questions of sample size.

LR intervals widen but match and posterior probability intervals do not? How does that work? How can we claim that uncertainty increases?

²⁷The calculations for the joint interval assume that because the worst- or best-case probability for one event is x and the worst- or best-case probability for another independent event is y , the worst- or best-case probability for their conjunction is xy . However, this conclusion does not follow if the margin of error (credible interval) is fixed. Just because the probability of an extreme value x for one variable X is .01, and so it is for the value y of another independent variable Y , it does not follow that the probability that those two independent variables take values x and y simultaneously is the same. In general, it is impossible to calculate the credible interval for the joint distribution based solely on the individual credible intervals corresponding to the individual events.

7 Evidence aggregation: the complex case

We looked at simple case of evidence aggregation: the pieces of evidence were about the same hypothesis and probabilistically independent. But different pieces of evidence can bear on different hypotheses and be probabilistically dependent. Think, for example, of two witnesses testifying in a trial about two related issues, say the defendant's whereabouts and the defendant's motive. In these more complex cases, precise probabilism handles evidence aggregation with the help of Bayesian networks. The graphical part of a Bayesian network consists of nodes and arrows. Arrows between nodes visually represents relationships of probabilistic dependence between different hypotheses and items of evidence, each corresponding to a node in the network. The numerical part of a Bayesian network describes the strengths of these dependencies. On a purely formal level, the numerical part consists of probability tables that are filled in with precise prior probabilities (for nodes without incoming arrows) or conditional probabilities (for nodes with incoming arrows).²⁸ Equipped with these input probabilities, the network can run the calculations about the output probabilities of interest.²⁹ We might be interested, for example, in the probability of a hypothesis given several items of evidence, while keeping track of dependencies between them. In the standard formulation, Bayesian networks run on precise probabilities, but can be extended to handle imprecise and higher order-probabilities. This is the topic of this section.

As an illustration, let us start with a Bayesian network developed by Fenton & Neil (2018). The network in Figure 7 represents the key items of evidence in the infamous British case *R. v. Clark* (EWCA Crim 54, 2000). Sally Clark, mother of two sons, witnessed her first son die in 1996 soon after birth. Her second son died in similar circumstances a few years later in 1998. These two consecutive deaths raised suspicion. One hypothesis about the cause of death is that Sally murdered her children. An alternative explanation is that both children died of Sudden Infant Death Syndrome (SIDS). At trial, however, the pediatrician Roy Meadow testified that the probability that a child from a family like the Clark's would die of SIDS was quite low, 1 in 8,543. Assuming probabilistic independence between the two events, the probability of both children dying of SIDS becomes extremely low. It equals the product of the two probabilities, approximately 1 in 73 million. Based on this low probability and signs of bruising on the bodies, Sally Clark was convicted of murder. The conviction was reversed on appeal thanks to new evidence, specifically, signs of a potentially lethal disease found in one of the bodies.

Why is a node for death missing in the network? That seems necessary as part of the evidence, no?

²⁸The simple case considered in the previous section would consist of a network with $k + 1$ nodes, with a root node for the hypothesis H and then k arrows going from H node to the evidence nodes E_1, E_2, \dots, E_k . The probability tables would be filled in with prior probabilities for H and conditional probabilities $P(E_i|H)$ and $P(E_i|\neg H)$, for any item of evidence E_i . Notice that these conditional probabilities are those occurring in the individual likelihood ratios $\frac{P(E_1|H)}{P(E_1|\neg H)}$. There is no need to rely on a Bayesian network in such a simple case because the dependencies between nodes are limited.

²⁹The calculations can quickly get out of hand, so softwares exist to perform the calculations automatically.

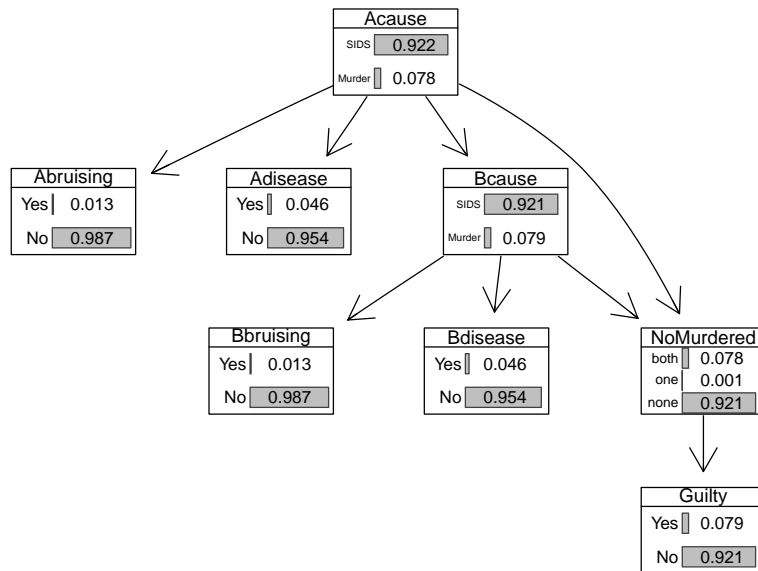


Figure 7: Bayesian network for the Sally Clark case, with marginal prior probabilities.

Much has been written about Sally Clark by philosophers and statisticians. The discussion has often focused on whether Meadow was allowed to assume, as he did, that the two SIDS death would be independent events. The assumption of independence delivered the low probability of 1 in 73 million by multiplying 1 in 8,543 by itself. Another, much discussed point was this: even if it was unlikely that two consecutive SIDS deaths would occur, it does not follow it was likely that Sally murdered her children.³⁰ A Bayesian network helps to avoid these mistakes. It also help to view the case holistically. The two consecutive deaths were an important piece of evidence, but other evidence was also important, including signs of bruising and signs of a lethal disease as they were discovered during the appeal process.

Unfortunately, Bayesian networks, in their standard formulation, inherit the shortcomings of precise probabilism. The choice of the input probabilities should be precise, and it is often unclear where the values come from or whether they are justified. Consider, for example, the probability that a death by SIDS would occur. How sure are we that this probability equals, exactly, 1 in 8,543? The figure Meadow used is a sample-based frequency. How big was that sample? How representative? Other input probabilities need to be entered in the network to carry out the calculations, for example, the probability that a mother would kill her son, or the probability that signs of bruising would be found given the hypothesis that Sally was trying to murder her child, and so on. There will be uncertainty about these probabilities, since they are based on sample frequencies or expert elicitation.

The standard response to these concerns is to invoke *sensitivity analysis*: a range of plausible values is tested. Say we are interested in the output probability that Sally is guilty. The network is updated by the known facts—the items of evidence—following standard Bayesian conditionalization. The input probabilities in the network are then assigned a range of possible values to see how they impact the output probability of Sally's guilt. Sensitivity analysis is another variant, perhaps more rudimentary, of the interval approach we considered earlier. In fact, Bayesian networks for reasoning with intervals and imprecise probabilities already

add references to imprecise BNs; see footnote.

³⁰One could reason that, since 1 in 73 million is a low probability, the alternative explanation, that Sally murdered her children, should be likely. But that a mother would do such a thing is also unlikely, perhaps even less likely than 1 in 73 million.

exist.³¹ But, as discussed earlier, imprecise probabilism, the interval approach and sensitivity analysis ignore the shape of the underlying distributions. They do not distinguish between probability measures (or point estimates) in terms of their plausibility, even though some will be more plausible than others. Moreover, if the sensitivity analysis is only guided by the values at the edges of the interval, these extremes will often play an undeservedly strong role.

These concerns can be addressed by recourse to higher-order probabilities. In a precise Bayesian network, each node is associated with a probability table determined by a finite list of numbers (precise probabilities). In an imprecise Bayesian network, each node is associated with a table determined by an interval of numbers. But suppose that, instead of precise numbers or intervals of numbers, we have distributions over the possible numbers to enter into the probability tables.³² An example of such higher-order Bayesian network for the Sally Clark case is represented in Figure 8.

The higher-order Bayesian network helps to investigate the impact of different items of evidence on Sally Clark's probability of guilt (see Figure 9). The starting point is the prior distribution for the Guilt node (first graph). Next, the network is updated with evidence showing signs of bruising on both children (second graph). Next, the assumption that both children lack signs of potentially lethal disease is added (third graph). Finally, we consider the state of the evidence at the time of the appellate case: signs of bruising existed on both children, but signs of lethal disease were discovered only on the first child. Interestingly, in the strongest case against Sally Clark (third graph), the median of the posterior distribution is above .95, but the uncertainty around that median is still quite wide.³³ This underscores the fact that relying on point estimates can lead to overconfidence.

³¹One can use uniform sampling with Bayesian networks to approximate the impreciser's commitments **Cite** <https://arxiv.org/abs/2302.09656>. Another approach is to rely on probabilistic programs with the restriction that the variables corresponding to probabilities are sampled from uniform distributions corresponding to the representor set. A critical survey of approaches along these lines shows that, in complex reasoning situations, "the imprecision of inferences increases rapidly as new premises are added to an argument".**add ref to** <https://www.sciencedirect.com/science/article/pii/S0004370296000215>.

³²The densities of interests can then be approximated by (1) sampling parameter values from the specified distributions, (2) plugging them into the construction of the BN, and (3) evaluating the probability of interest in that precise BN. The list of the probabilities thus obtained will approximate the density of interest. In what follows we will work with sample sizes of 10k.

³³The lower limit of the 89% Highest Posterior Density Intervals (HPDI) is at .83.

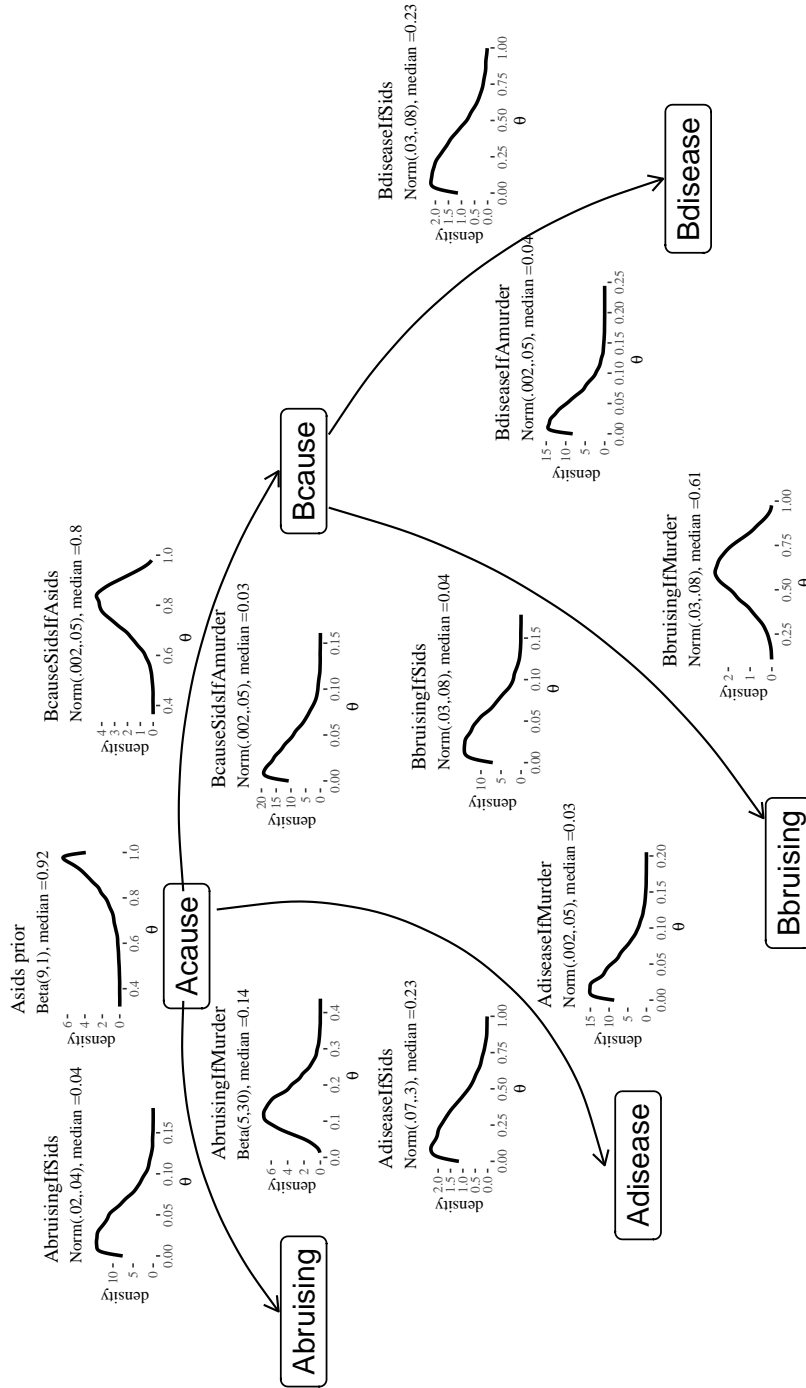


Figure 8: An illustration of a probabilistic program for the Sally Clark case.

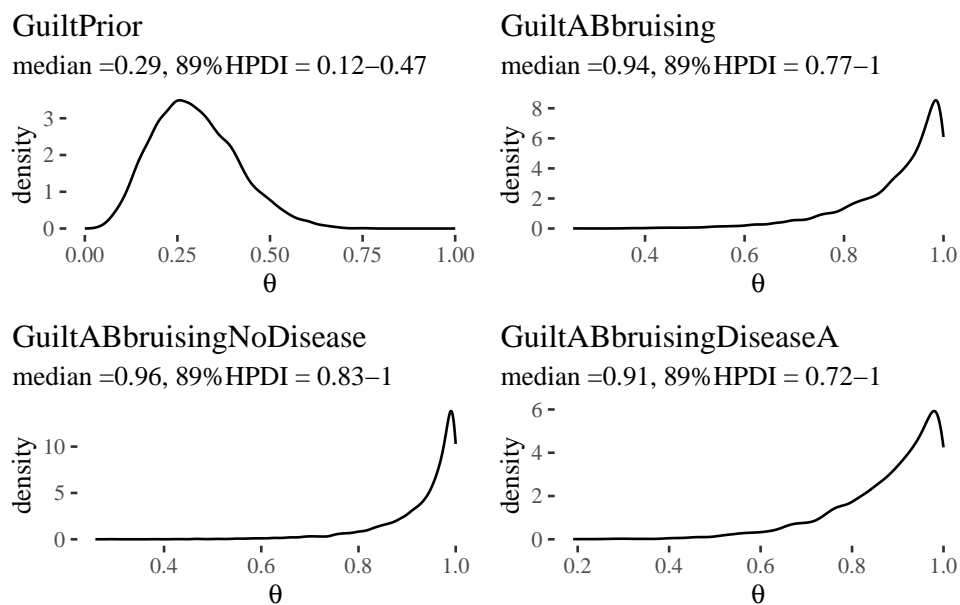


Figure 9: Impact of incoming evidence in the Sally Clark case.

One question that arises is how this approach relates to the standard method of using likelihood ratios to report the value of the evidence. On this approach, the conditional probabilities that are used in the likelihood ratio calculations are estimated and come in a package with an uncertainty about them. Accordingly, these uncertainties propagate: to estimate the likelihood ratio while keeping track of the uncertainty involved, we can sample probabilities from the selected distributions appropriate for the conditional probabilities needed for the calculations, then divide the corresponding samples, obtaining a sample of likelihood ratios, thus approximating the density capturing the recommended uncertainty about the likelihood ratio. Uncertainty about likelihood ratio is just propagated uncertainty about the involved conditional probabilities. For instance, we can use this tool to gauge our uncertainty about the likelihood ratios corresponding to the signs of bruising in son A and the presence of the symptoms of a potentially lethal disease in son A (Figure 10).

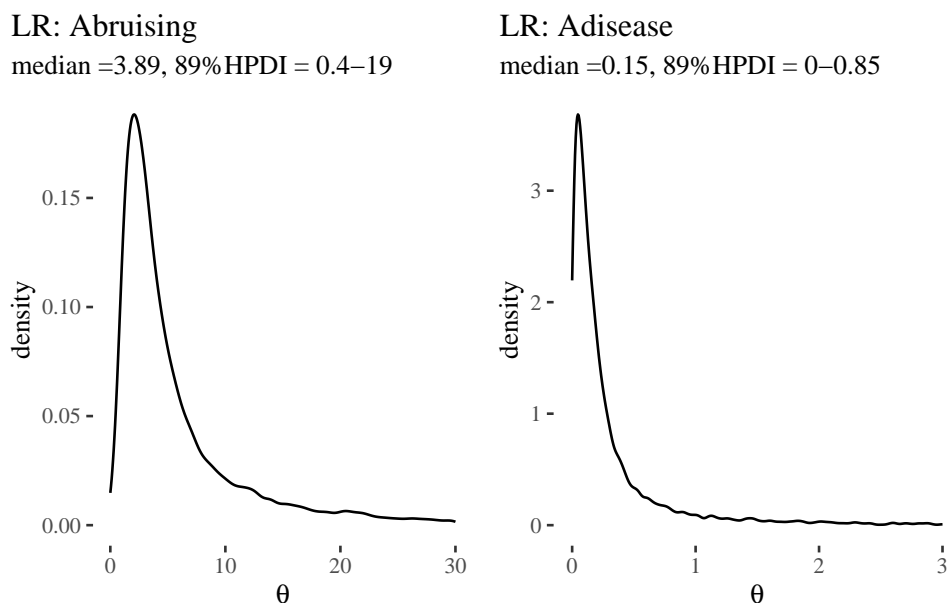


Figure 10: Likelihood ratios for bruising and signs of disease in child A in the Sally Clark case.

8 Conclusion

We have argued that higher-order probabilism outperforms both precise and imprecise probabilism. It is able to model scenarios that the other two cannot model, for example, the case of uneven bias. In addition, higher-order probabilism does not fall prey to difficulties peculiar to imprecise probabilism, such as belief inertia and the lack of proper scoring rules. We have also identified a novel set of problems for precise and imprecise probabilism, mostly stemming from how to evaluate, in the aggregate, multiple pieces of evidence. Here again, higher-order probabilism fares better.

Some might dislike the idea of going higher-order for a number of reasons, for example, unnecessary complexity. This is a line taken by Bradley, who refuses to go higher-order for the following reason:

Why is sets of probabilities the right level to stop the regress at? Why not sets of sets? Why not second-order probabilities? Why not single probability functions? This is something of a pragmatic choice. The further we allow this regress to continue, the harder it is to deal with these belief representing objects. So let's not go further than we need. 131-132

We have shown that given the difficulties of precise and imprecise probabilism, we are not going further than we need in using higher-order probabilities. The pragmatic concerns one might have are at best unclear.

We should underscore that, mathematically, we do not propose anything radically new. Concepts from the Bayesian toolkit that can model higher-order uncertainty already exist. Our suggestion is that they have been under-appreciated in formal epistemology and should be more widely used. This is not to say that there is no need for any novel technical work. For example, we still need a proper accuracy argument in defense of higher-order probabilism. Will an agent who relies on higher-order probabilities would accuracy-dominate one who relies on just first-order probabilistics? We leave this as an open question. Another concern is the lack of a clear semantics for higher-order probabilities. While a more elaborate account is beyond the scope of this paper, the answer should gesture at a modification of the framework

of probabilistic frames (Dorst, 2022b, 2022a). Start with a set of possible worlds W . Suppose you consider a class of probability distributions D , a finite list of atomic sentences q_1, \dots, q_2 corresponding to subsets of W , and a selection of true probability hypotheses C (think of the latter as omniscient distributions, $C \subseteq D$, but in principle this restriction can be dropped if need be). Each possible world $w \in W$ and a proposition $p \subseteq W$ come with their true probability distribution, $C_{w,p} \in D$ corresponding to the true probability of p in w , and the distribution that the expert assigns to p in w , $P_{w,p} \in D$. Then, various propositions involving distributions can be seen as sets of possible worlds, for instance, the proposition that the expert assigns d to p is the set of worlds w such that $P_{w,p} = d$.³⁴

Appendix: the strict propriety of I_{kl}

The fact that I_{KL} is strictly proper as applied to second-order probabilities is not very surprising. However, in the existing literature, the proof is not usually explicitly given, and some of the pieces are not present in philosophical literature. So we tried to include the whole chain of thought, warning that some of these results are already known and all we did was making the proofs more presentable, and pointing out new elements in the reasoning. Let us start with a definition of concavity.

Definition 1 (concavity). *A function f is convex over an interval (a, b) just in case for all $x_1, x_2 \in (a, b)$ and $0 \leq \lambda \leq 1$ we have:*

$$f(\lambda x_1 + (1 - \lambda)x_2) \leq \lambda f(x_1) + (1 - \lambda)f(x_2)$$

A function f is concave just in case:

$$f(\lambda x_1 + (1 - \lambda)x_2) \geq \lambda f(x_1) + (1 - \lambda)f(x_2)$$

A function f is strictly concave just in case the equality holds only if either $\lambda = 0$ or $\lambda = 1$.

For us it is important that if a function is twice differentiable on an interval, then it is (strictly) concave just in case its second derivative is non-positive (negative). In particular, as $(\log_2(x))'' = -\frac{1}{x^2 \ln(2)}$, \log_2 is strictly concave over its domain.³⁵

NL: Why ""?

Lemma 1 (Jensen's inequality). *If f is concave, and g is any function of a random variable, $\mathbb{E}(f(g(x))) \leq f(\mathbb{E}(g(x)))$. If f is strictly concave, the equality holds only if $g(x) = \mathbb{E}g(x)$, that is, if $g(x)$ is constant everywhere.*

Proof. For the base case consider a two-point mass probability function. Then,

$$p_1 f(g(x_1)) + p_2 f(g(x_2)) \leq f(p_1 g(x_1) + p_2 g(x_2))$$

follows directly from the definition of concavity, if we take $\lambda = p_1$, $(1 - \lambda) = p_2$, and substitute $g(x_1)$ and $g(x_2)$ for x_1 and x_2 .

Now, suppose that $p_1 f(g(x_1)) + p_2 f(g(x_2)) = f(p_1 g(x_1) + p_2 g(x_2))$ and that f is strictly concave. That means either $(p_1 = 1 \wedge p_2 = 0)$, or $(p_1 = 0 \wedge p_2 = 1)$. Then either x always takes value x_1 , in the former case, or always takes value x_2 , in the latter case. $\mathbb{E}g(x) = p_1 g(x_1) + p_2 g(x_2)$, which equals $g(x_1)$ in the former case and $g(x_2)$ in the latter.

³⁴There is at least one important difference between this approach and that developed by Dorst. His framework is untyped, which allows for an enlightening discussion of the principle of reflection and alternatives to it. In this paper, we prefer to keep this complexity apart and use an explicitly typed set-up.

³⁵I line with the rest of the paper, we'll work with log base 2. We could equally well use any other basis.

Now suppose Jensen's inequality and the consequence of strict convexity holds for $k - 1$ mass points. Write $p'_i = \frac{p_i}{1-p_k}$ for $i = 1, 2, \dots, k - 1$. We now reason:

$$\begin{aligned}
\sum_{i=1}^k p_i f(g(x_i)) &= p_k f(g(x_k)) + (1 - p_k) \sum_{i=1}^{k-1} p'_i f(g(x_i)) \\
&\leq p_k f(g(x_k)) + (1 - p_k) f\left(\sum_{i=1}^{k-1} p'_i g(x_i)\right) && \text{by the induction hypothesis} \\
&\leq f\left(p_k g(x_k) + (1 - p_k) \sum_{i=1}^{k-1} p'_i g(x_i)\right) && \text{by the base case} \\
&= f\left(\sum_{i=1}^k p_i g(x_i)\right)
\end{aligned}$$

Notice also that at the induction hypothesis application stage we know that the equality holds only if $p_k = 1 \vee p + k = 0$. In the former case $g(x)$ always takes value $x_k = \mathbb{E}g(x)$. In the latter case, p_k can be safely ignored and $\sum_{i=1}^k p_i g(x_i) = \sum_{i=1}^{k-1} p'_i g(x_i)$ and by the induction hypothesis we already know that $\mathbb{E}g(x) = g(x)$. □

In particular, the claim holds if we take $g(x)$ to be $\frac{q(x)}{p(x)}$ (were both p and q are probability mass functions), and f to be \log_2 . Then, given that A is the support set of p , we have:

$$\sum_{x \in A} p(x) \log_2 \frac{q(x)}{p(x)} \leq \log_2 \sum_{x \in A} p(x) \frac{q(x)}{p(x)}$$

Moreover, the equality holds only if $\frac{q(x)}{p(x)}$ is constant, that is, only if p and q are the same pmfs. Let's use this in the proof of the following lemma.

Lemma 2 (Information inequality). *For two probability mass functions p, q , $D_{\text{KL}}(p, q) \geq 0$ with equality iff $p = q$.*

Proof. Let A be the support set of p , and let q be a probability mass function whose support

is B .

$$\begin{aligned}
-D_{\text{KL}}(p, q) &= -\sum_{x \in A} p(x) \log_2 \frac{p(x)}{q(x)} && \text{(by definition)} \\
&= \sum_{x \in A} p(x) - (\log_2 p(x) - \log_2 q(x)) \\
&= \sum_{x \in A} p(x) (\log_2 q(x) - \log_2 p(x)) \\
&= \sum_{x \in A} p(x) \log_2 \frac{q(x)}{p(x)} \\
&\leq \log_2 \sum_{x \in A} p(x) \frac{q(x)}{p(x)} && \text{by Jensen's inequality} \\
&\text{(and the equality holds only if } p = q) \\
&= \log_2 \sum_{x \in A} q(x) \\
&\leq \log_2 \sum_{x \in B} q(x) \\
&= \log(1) = 0
\end{aligned}$$

□

Observe now that D_{KL} can be decomposed in terms of cross-entropy and entropy.

Lemma 3 (decomposition). $D_{\text{KL}} = H(p, q) - H(p)$.

Proof.

$$\begin{aligned}
D_{\text{KL}}(p, q) &= \sum_{p_i} (\log_2 p_i - \log_2 q_i) \\
&= -\sum_{p_i} (\log_2 q_i - \log_2 p_i) \\
&= -\sum_{p_i} \log_2 q_i - \sum_{p_i} -\log_2 p_i \\
&= \underbrace{-\sum_{p_i} \log_2 q_i}_{H(p, q)} - \underbrace{\sum_{p_i} -\log_2 p_i}_{H(p)}
\end{aligned}$$

□

With information inequality this easily entails Gibbs' inequality:

Lemma 4 (Gibbs' inequality). $H(p, q) \geq H(p)$ with identity only if $p = q$.

We are done with our theoretical set-up, which is already common knowledge, except presented in an orderly manner in one place. Now we present our argument for the claim that the above entails the propriety of I_{KL} . First, let's systematize the notation. Consider a discretization of the parameter space $[0, 1]$ into n equally spaced values $\theta_1, \dots, \theta_n$. For each i the

“true” second-order distribution if the true parameter indeed is θ_i —we’ll call it the indicator of θ_i — which is defined by

$$Ind^k(\theta_i) = \begin{cases} 1 & \text{if } \theta_i = \theta_k \\ 0 & \text{otherwise} \end{cases}$$

We will write Ind_i^k instead of $Ind^k(\theta_i)$.

Now consider a probability distribution p over this parameter space, assigning probabilities p_1, \dots, p_n to $\theta_1, \dots, \theta_n$ respectively. It is to be evaluated in terms of inaccuracy from the perspective of a given ‘true’ value θ_k . The inaccuracy of p if θ_k is the ‘true’ value, is the divergence between Ind^k and p .

$$\begin{aligned} I_{KL}(p, \theta_k) &= D_{KL}(Ind^k || p) \\ &= \sum_{i=1}^n Ind_i^k (\log_2 Ind_i^k - \log_2 p_i) \end{aligned}$$

Note now that for $j \neq k$ we have $Ind_j^k = 0$ and so $Ind_j^k (\log_2 Ind_j^k - \log_2 p_j) = 0$. Therefore we continue:

$$= Ind_k^k (\log_2 Ind_k^k - \log_2 p_k)$$

Further, $Ind_k^k = 1$ and therefore $\log_2 Ind_k^k = 0$, so we simplify:

$$= -\log_2 p_k$$

Now, let’s think about expected values. First, what is the inaccuracy of p as expected by p , $\mathbb{E}I_{DK}(p, p)$?

$$\begin{aligned} \mathbb{E}I_{DK}(p, p) &= \sum_{i=1}^n p_i I_{DK}(p, \theta_i) \\ &= \sum_{i=1}^n p_i (-\log_2 p_i) \\ &= -\sum_{i=1}^n p_i \log_2 p_i = H(p) \end{aligned}$$

Analogously, the inaccuracy of q as expected from the perspective of p is:

$$\begin{aligned} \mathbb{E}I_{DK}(p, q) &= \sum_{i=1}^n p_i (-\log_2 q_i) \\ &= -\sum_{i=1}^n p_i \log_2 q_i = H(p, q) \end{aligned}$$

But that means, by Gibbs’ inequality, that $\mathbb{E}I_{DK}(p, q) \geq \mathbb{E}I_{DK}(p, p)$ unless $p = q$, which completes the proof.

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