

A probabilistic diagnostic tool to assess Laplace approximations: proof of concept and non-asymptotic experimentation

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Abstract

In many statistical models, we need to integrate functions that may be high-dimensional. Such integrals may be impossible to compute exactly, or too expensive to compute numerically. Instead, we can use the *Laplace approximation* for the integral. This approximation is exact if the function is proportional to the density of a normal distribution; therefore, its effectiveness may depend intimately on the true shape of the function. To assess the quality of the approximation, we use *probabilistic numerics*: recasting the approximation problem in the framework of probability theory. In this probabilistic approach, uncertainty and variability don't come from a frequentist notion of randomness, but rather from the fact that the function may only be partially known. We use this framework to develop a diagnostic tool for the Laplace approximation, modelling the function and its integral as a Gaussian process and devising a “test” by conditioning on a finite number of function values. We will discuss approaches for designing and optimizing such a tool and demonstrate it on known sample functions, highlighting in particular the challenges one may face in high dimensions.

1 Introduction

Coming soon. Some combination of abstract (above) and framework (below). Specifically mention:

1. Use cases including state space models
2. That we are building on the work of Zhou [2]
3. That this is non-asymptotic and not intended as a substitute for full-on MC integration or BQ - rather as a “middle-ground” amount of effort.
4. The goal is to “test” the assumptions underlying the Laplace approximation (e.g. “how Gaussian is this function?”). The Laplace approximation

may still hold for a non-Gaussian shape, but such a function should be rejected by our diagnostic (“sufficiently non-Gaussian things warrant further attention”), at which point a more involved integration would show that the approximation was fine after all.

2 Framework and notation

Consider a positive function $f : \mathbb{R}^d \rightarrow \mathbb{R}_{>0}$. The object of interest for the diagnostic is the integral $F = \int_{\mathbb{R}^d} f(t) dt$. In practical applications [citation needed], typically f and F are actually functions with an additional argument vector of structural parameters θ , with $t \in \mathbb{R}^d$ a vector of nuisance parameters to be marginalized. For instance, f may be a joint probability density for (θ, t) , in which case F would be the marginal distribution of θ after integrating over t . To reflect this common setting, Zhou [2] called f and F the *full* and *target* functions, respectively. For the present discussion, non-marginalized arguments θ are not relevant, so any dependence on them is omitted and f and F are simply called the *true* function and integral, respectively.

Suppose now that f has all second-order partial derivatives¹, and a (local) maximum at some point $\hat{t} \in \mathbb{R}^d$. To reflect the use case where f is a density, \hat{t} is called a *mode*. Let H be the Hessian of $\log f$ at \hat{t} , and suppose that it is negative-definite (i.e. that f is log-concave at the mode). The first step in arriving at the Laplace approximation [citation needed] for F is to take a second-order Taylor expansion of $\log f$ about \hat{t} . Noting that all first-order partial derivatives of $\log f$ are equal to zero at the mode, this approximation is

$$\log f(t) \approx \log f(\hat{t}) + \frac{1}{2} (t - \hat{t})^\top H (t - \hat{t}). \quad (1)$$

Exponentiating the right side of (1) gives an approximation for f in the form of (up to normalizing constants) a Gaussian density centered at \hat{t} with covariance matrix $-H^{-1}$. In turn, integrating this exponentiated function produces the *Laplace approximation*²

$$\begin{aligned} F \approx L(f) &:= f(\hat{t}) \int_{\mathbb{R}^d} \exp \left[\frac{1}{2} (t - \hat{t})^\top H (t - \hat{t}) \right] dt \\ &= f(\hat{t}) \sqrt{(2\pi)^d \det(-H^{-1})}. \end{aligned} \quad (2)$$

The Laplace approximation is exact (or “true”) if f is itself proportional to a Gaussian density. There are other function shapes for which this may be the case, but such instances may be thought of as “coincidence”. Certainly, the construction of the Laplace approximation via (1) is based on an assumption of approximately Gaussian shape, and this is assumption is our primary interest in developing a diagnostic.

¹TODO: check actual assumptions for Laplace. Are third derivatives necessary?

²TODO: get citation for this. In particular, there are a couple of variations I’ve seen in the form for “Laplace’s method” in the Bayesian literature. For instance, sometimes $\log f$ is multiplied by a constant (usually sample size) before exponentiating.

3 Probabilistic numerics and Bayesian quadrature

³Broadly speaking, probabilistic numerics is the use of probability theory, from a somewhat Bayesian perspective, to simultaneously perform estimation and uncertainty quantification in standard numerical problems [citation needed]. For instance, Chkrebtii et al. [1] developed a probabilistic solver for differential equations. For a given equation, they jointly modelled the function and its derivatives with a Gaussian process prior, then sequentially conditioned on true derivative values to conduct posterior inference on the entire solution.

The approach briefly described above - using Gaussian process priors and finitely many function values to obtain posteriors for the functions and quantities of interest - is at the core of many probabilistic numerical methods. In particular, it is the standard framework with which *Bayesian quadrature* (BQ) is usually conducted. [COMING SOON: citations and context for BQ. As ‘‘original’’ as possible]

The machinery of BQ can be used to develop a probabilistic diagnostic for the Laplace approximation, as in [2]. Recalling the notation of Section 2, f is modelled with a Gaussian process prior. The mean function of the GP prior, m_0^t , is taken to be the Gaussian function underpinning (1) and (2):

$$m_0^t(t) := f(\hat{t}) \exp \left[\frac{1}{2} (t - \hat{t})^\top H (t - \hat{t}) \right], t \in \mathbb{R}^d. \quad (3)$$

The covariance operator for the GP is a (positive-definite) kernel C_0^t on $\mathbb{R}^d \times \mathbb{R}^d$ to be defined later.

By the projection property of Gaussian processes [citation/clarification needed - will fill in later], such a prior on f induces a scalar Normal prior on F with mean $m_0 := \int_{\mathbb{R}^d} m_0^t(t) dt = L(f)$ and variance $C_0 := \int_{\mathbb{R}^d} \int_{\mathbb{R}^d} C_0^t(t, u) dt du$, provided all relevant quantities exist and are finite.⁴

In what follows, let $\mathbf{s} = (s_1, \dots, s_n)^\top \in \mathbb{R}^{d \times n}$ be a row-wise concatenation of n vectors in \mathbb{R}^d . Then, for instance, the notation $f(\mathbf{s})$ will refer to the column vector $(f(s_1), \dots, f(s_n))^\top \in \mathbb{R}^n$, and $C_0^t(\mathbf{s}, \mathbf{s})$ will denote the $n \times n$ matrix with $(i, j)^{\text{th}}$ entry $C_0^t(s_i, s_j)$. As in [2]⁵, one may use true function values at the *interrogation points* \mathbf{s} to obtain a posterior distribution for f (with a slight abuse of notation):

$$f \mid [f(\mathbf{s})] \sim \mathcal{GP}(m_1^t, C_t^1), \quad (4)$$

$$m_1^t(t) = m_0^t(t) + C_t^0(t, \mathbf{s})^\top [C_t^0(\mathbf{s}, \mathbf{s})]^{-1} (f(\mathbf{s}) - m_0^t(\mathbf{s})), \quad (5)$$

$$C_1^t(t, u) = C_0^t(t, u) - C_t^0(t, \mathbf{s})^\top [C_t^0(\mathbf{s}, \mathbf{s})]^{-1} C_t^0(u, \mathbf{s}). \quad (6)$$

³TODO: pad out the introductory PN/BQ stuff.

⁴TODO: check this and find citations. Need to check the conditions under which it holds on an infinite domain and state those more precisely

⁵And maybe another PN/BQ citation, since this is the standard thing to do. A citation for the posterior update of a GP may be good too.

In turn, the posterior distribution on the integral F is

$$F \mid [f(\mathbf{s})] \sim \mathcal{N}(m_1, C_1), \quad (7)$$

$$m_1 = L(f) + \left[\int_{\mathbb{R}^d} C_t^0(t, \mathbf{s}) d\mathbf{z} \right]^\top [C_t^0(\mathbf{s}, \mathbf{s})]^{-1} (f(\mathbf{s}) - m_0^t(\mathbf{s})), \quad (8)$$

$$C_1 = C_0 - \left[\int_{\mathbb{R}^d} C_t^0(t, \mathbf{s}) d\mathbf{z} \right]^\top [C_t^0(\mathbf{s}, \mathbf{s})]^{-1} \left[\int_{\mathbb{R}^d} C_t^0(t, \mathbf{s}) d\mathbf{z} \right], \quad (9)$$

where the integrals are row-wise over \mathbf{s} :

$$\int_{\mathbb{R}^d} C_t^0(t, \mathbf{s}) d\mathbf{z} = \left(\int_{\mathbb{R}^d} C_t^0(t, s_1) d\mathbf{z}, \dots, \int_{\mathbb{R}^d} C_t^0(t, s_n) d\mathbf{z} \right)^\top.$$

The posterior (7) will serve as the diagnostic for the Laplace approximation. Borrowing from the traditional notion of hypothesis testing, one may deem the Laplace approximation (or perhaps more accurately, the shape assumptions motivating it) acceptable or valid if $L(f)$ falls within the range spanned by the $(0.025, 0.975)$ quantiles of (7) (the 95% “confidence interval” centered at the posterior mean). Conversely, if $L(f)$ is outside of this interval, the Laplace approximation would be deemed inappropriate, and one would proceed to use a more involved method to estimate F . Traditionally [add old BQ citation], the goal of BQ is convergence to the true integral: choosing the covariance kernel and interrogation points such that (8) and (9) are close to F and 0, respectively. This is not our main goal in designing the diagnostic, which is intended to be decidedly non-asymptotic: rather, it should be able to effectively facilitate the aforementioned “hypothesis” test with as little computational cost as possible, whether or not that results in a good integral estimate.

4 Placement of interrogation points

Typically, the number of points n required to estimate a d -dimensional integral to within some error tolerance increases exponentially in d [citation needed, especially for BQ]. This creates an unfortunate computational bottleneck in BQ, as the main cost is inverting the $n \times n$ matrix $C_0^t(\mathbf{s}, \mathbf{s})$. However, the goal of this diagnostic is to efficiently test the Gaussian shape assumption underpinning the Laplace approximation, with accurate integral estimation as an afterthought. Presumably this goal can be achieved with fewer interrogation points than a full BQ, allowing in principle for easier scaling to high dimensions.

First, assume without loss of generality that \hat{t} is at the origin and $H = -I$. This ensures that the Gaussian approximation to f , m_0^t , is proportional to a standard Normal density. If this not the case, recall that H is negative-definite, so there exists a matrix G such that $G^\top H G = -I$. For instance, $G = V [\sqrt{-D}]^{-1}$ from the eigendecomposition $H = V D V^\top$ serves this purpose. Then the aforementioned assumptions may be enforced by replacing f with the function $t \mapsto f(Gt + \hat{t})$.

As it pertains to the selection of interrogation points \mathbf{s} , this transformation serves two purposes. The first is to “rotate” the domain of f so that its direction of strongest curvature corresponds to one of the axes⁶. Heuristically, this means that the values of f along the axes will offer the most pertinent information about its shape. The second purpose is scaling and shifting, which allows interrogation points to be defined in a very intuitive way. This is best explained with an example: in two dimensions, when $H = -I$ and $\hat{t} = (0, 0)$, an interrogation point at $(m, 0)$ corresponds to a point that is m “standard deviations” (of the bivariate Normal distribution with density proportional to m_0^t) from the mode along the x -axis. With these ideas in mind, we propose to use a d -dimensional “cross-shaped” grid of interrogation points⁷ consisting of the mode and additional points placed at regular spacings along each axis. Such grids will be characterized by the distances between consecutive points along the axes and the distance between the mode and the extremal points, both in terms of “standard deviations of m_0^t ” as described above. For instance, one may wish to place interrogation points at half-integer multiples of the standard deviation, up to a maximum of three standard deviations, along each axis. This corresponds to points of the form $\pm \frac{m}{2} e_i$ where $m = 0, 1, \dots, 6$ and e_i is a standard basis vector in \mathbb{R}^d , $i = 1, \dots, d$. Although the use of these ideas requires a somewhat costly eigendecomposition of H , we believe that the alignment of shape information with the axes outweighs any such costs: the proposed cross-shaped grids grow linearly in size with d , bypassing much of the computational cost associated with more involved quadrature techniques. For instance, the grid given as an example above consists of only $n = 12d + 1$ points, and is expected to convey enough shape information to make the diagnostic work.

5 A finite-integral covariance kernel

The choice is a covariance kernel is important in determining the behaviour of a probabilistic quadrature method. Chkrebtii et al. [1], and subsequently Zhou [2], used a self-convolution of the popular squared exponential kernel [sources/further info on sq exp kernel could be inserted if needed]:

$$C_0^t(t, u) = \left(\frac{\sqrt{\pi}\lambda}{\alpha} \right)^d \exp \left[-\frac{\|t - u\|^2}{4\lambda^2} \right], \quad (10)$$

where, respectively, the *length-scale* and *precision* hyperparameters λ and α control the sample smoothness and spread of the GP.

A problem arises if one wishes to use this kernel without modification: its integral over $\mathbb{R}^d \times \mathbb{R}^d$ diverges, so the prior distribution assigned to F will

⁶This point can be made clear with some linear algebra and multivariate calculus. First note that the second directional derivative of $\log f$ at the mode is always negative and is minimized along the direction of some eigenvector of H . Finally observe that the Hessian of f at the mode has the same eigenvectors as H , and the rotational part of the transformation maps them to standard basis vectors.

⁷If I recall correctly, somebody made this suggestion to Dave at a conference in early 2018.

have infinite variance. Some practitioners avoid this problem by integrating over finite regions rather than the whole of \mathbb{R}^d : Chkrebtii et al. [1] considered ODE’s defined on compact intervals, and Zhou [2] took integrals over a region bounded by the extremal interrogation points. It is perhaps more common in BQ literature **[todo: add citations]** to integrate with respect to a probability measure Π on \mathbb{R}^d . In that case, the object of interest is $\int_{\mathbb{R}^d} f(t) d\Pi(t)$, for which the prior mean and variance are defined by, respectively, the integral of m_0^t w.r.t Π and the integral of C_0^t w.r.t. the product measure $\Pi \times \Pi$. With this framework, all of the necessary integrals converge, in contrast to our setting where integrals are taken w.r.t. the Lebesgue measure.

To solve the problem of infinite variance, we take a different approach⁸: adding a “decay” factor to the covariance kernel so that it has finite integral over $\mathbb{R}^d \times \mathbb{R}^d$. The modified kernel used through this manuscript is

$$C_0^t(t, u) = f(\hat{t})^2 \left(\frac{\sqrt{\pi}\lambda}{\alpha} \right)^d \exp \left[-\frac{\|t - u\|^2}{4\lambda^2} \right] \exp \left[-\frac{\|t\|^2 + \|u\|^2}{4\gamma^2} \right]. \quad (11)$$

The new hyperparameter γ controls the rate at which the GP prior variance, $C_0^t(t, t)$, decays as t moves away from the origin. Intuitively it is reasonable to impose such behaviour on the GP prior: any function f to which this diagnostic applies would certainly decay to 0 in the tails, so the prior assumption that uncertainty/variability in its values decreases with distance from the mode (assumed to be at the origin, as per Section 4) is appropriate⁹. With this modification, the covariance kernel (11) is proportional to a $2d$ -dimensional Gaussian density. Figure 1 shows a visual comparison between this kernel and the one given by (10) in the one-dimensional case. Note that (10) is the limiting case of (11) as $\gamma \rightarrow \infty$. Using this modified kernel with $\gamma \in (0, \infty)$ ensures that the scalar Normal prior on F will have finite variance **[todo: make sure all the GP convergence stuff is actually valid for this]** given by

$$C_0 = f(\hat{t})^2 \left[\frac{4\gamma^2\lambda^2\sqrt{\pi^3}}{\alpha\sqrt{2\gamma^2 + \lambda^2}} \right]^d. \quad (12)$$

The factor of $f(\hat{t})^2$ ensures scale invariance in the diagnostic. The posterior variance of F (9) depends only on the placement of interrogation points **[this is well-known in BQ, so could maybe provide source]**. In turn, interrogation point locations are indirectly based on the shape of $\log f$ as described in Section 4. Because the shape of $\log(Mf)$ is the same for any scaling constant $M > 0$, the posterior variance will be unchanged by such scaling even though the difference between the posterior and prior means is $M(m_1 - L(f))$. Therefore, without explicitly incorporating scale information into the covariance, the rejection criteria for the diagnostic will be scale-dependent: an undesirable trait given that the shape of the true function and the *proportional* difference

⁸Should Richard be a last author? Modifying the covariance kernel to decay was his idea (not specifically *how* to modify it, though)

⁹This was also a point made by Richard originally.

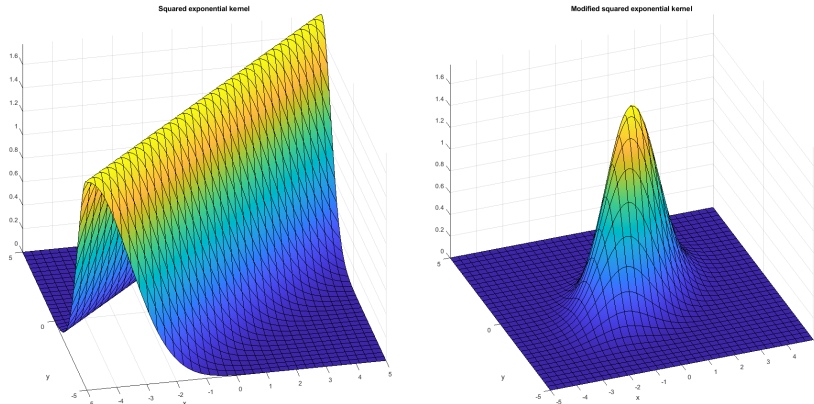


Figure 1: For $d = 1$, the usual squared exponential kernel with $\lambda = \alpha = 1$ (left) vs. the modified decaying version with $\lambda = \alpha = \gamma = 1$ (right).

between its integral and the Laplace approximation do not depend on scale¹⁰. Incorporating the factor of $f(\hat{t})^2$ into the covariance kernel ensures that the Laplace approximation is rejected for f iff it is rejected for Mf for all $M > 0$.

6 Calibration and Experiments

Literature on BQ contains several approaches to optimizing both interrogation point placements and covariance hyperparameters [sources and examples coming soon]. However, these approaches are usually aimed towards a high-accuracy, low-variance estimate of the integral, which differs from the goal of the diagnostic. In particular, the diagnostic should *fail* to reject any function whose shape is sufficiently close to Gaussian for the Laplace approximation to be reasonable. Thus, our approach to optimization will be based on somewhat heuristic calibrations.

Let $T_{\nu,d}$ denote the density of the d -variate T distribution with ν degrees of freedom. Such a density has heavier tails than a d -dimensional Gaussian density, and so its integral is underestimated by the Laplace approximation. However, the Gaussian is the limiting case of the T density as $\nu \rightarrow \infty$. Therefore, for some large value of ν , the shape of $T_{\nu,d}$ is “sufficiently Gaussian” and the diagnostic should be calibrated *not* to reject its Laplace approximation, which will be close to the true integral value of 1. We will expand on this shortly, but first it is useful for exploratory purposes to see how the posterior mean (8) and variance (9) depend on the hyperparameters of the covariance kernel. For this preliminary

¹⁰Another thing briefly mentioned by Richard and only recently remembered/implemented by me.

visualization, we use true functions of $T_{1,1}$ and $T_{1,2}$ (for which the tails are heavy enough to warrant a definite rejection of the Laplace approximation), a grid of interrogation points placed at $0, \pm 1, \pm 2$ standard deviations along each axis, and $\alpha = 1$. The latter choice is made because the only effect of α is on the scale of the posterior variance, and so we are more interested in the effect of λ and γ on diagnostic behaviour.

Figure 2 shows variation in the posterior mean and variance over a range of λ and γ values. Jagged edges in the plots at high λ 's are likely indicative of numerical instability due to oversmoothing, which pushes the matrix $C_0^t(\mathbf{s}, \mathbf{s})$ towards singularity. The posterior reduction in variance increases with λ , but evidently to a lesser extent than the prior variance, so that the posterior variance is ultimately higher for large λ . The effect of λ will be explored later in more detail, but the effect of γ is interesting enough to warrant discussion here. Although γ appears to influence the variance “correction” (the difference between prior and posterior variance) and posterior mean at low values (and causes some possible numerical instability for the latter), its effect levels off considerably beyond a certain threshold. Indeed, when γ is sufficiently large, its primary contribution is towards the scale of the prior variance (12). The behaviour is similar for the 2-dimensional case, as shown in Figure 3. The main difference is that the posterior variance increases more dramatically with both γ and λ , which is to be expected due to the prior variance being $\mathcal{O}(\lambda^{2d})$ and $\mathcal{O}(\gamma^{2d})$. There also appears to be a greater range of variability with respect to λ in the posterior mean, and perhaps in the variance correction as well.

Let us now return to the issue of calibration. Recall that the limiting behaviour of $T_{d,\nu}$ with respect to ν makes it close to a Gaussian shape for high degrees of freedom. This ensures that the Laplace approximation,

$$L(T_{\nu,d}) = \left(\frac{2}{\nu + d} \right)^{\frac{d}{2}} \frac{\Gamma\left(\frac{\nu+d}{2}\right)}{\Gamma\left(\frac{\nu}{2}\right)}, \quad (13)$$

is increasing in ν and approaches 1 as $\nu \rightarrow \infty$. For a given dimension d , let ν_d be the smallest integer such that $L(T_{\nu_d,d}) \geq 0.95$. Densities with more than ν_d degrees of freedom are close enough in shape to Gaussians that their Laplace approximations are within 5% of the true integral value; conversely, those with lower degrees of freedom have heavier tails and Laplace approximations that underestimate the true integral by over 5%. Thus, $T_{\nu_d,d}$ will serve as the “edge case” function used to calibrate the diagnostic. Specifically, for a given dimension d and grid of interrogation points \mathbf{s} ,

1. α should be set so that $L(T_{\nu_d,d})$ is on the boundary of the non-rejection region (the interval spanned by the (0.025, 0.975) quantiles of the posterior $\mathcal{N}(m_1, C_1)$ distribution); and
2. λ and γ should be set to minimize discrepancy between the true function $T_{\nu_d,d}$ and the posterior mean function m_1^t (5).

The first requirement ensures that $T_{\nu_d,d}$ acts as a sort of “threshold”, as intended. With α set in this way, $T_{\nu_d,d}$ is “just Gaussian enough” for the Laplace

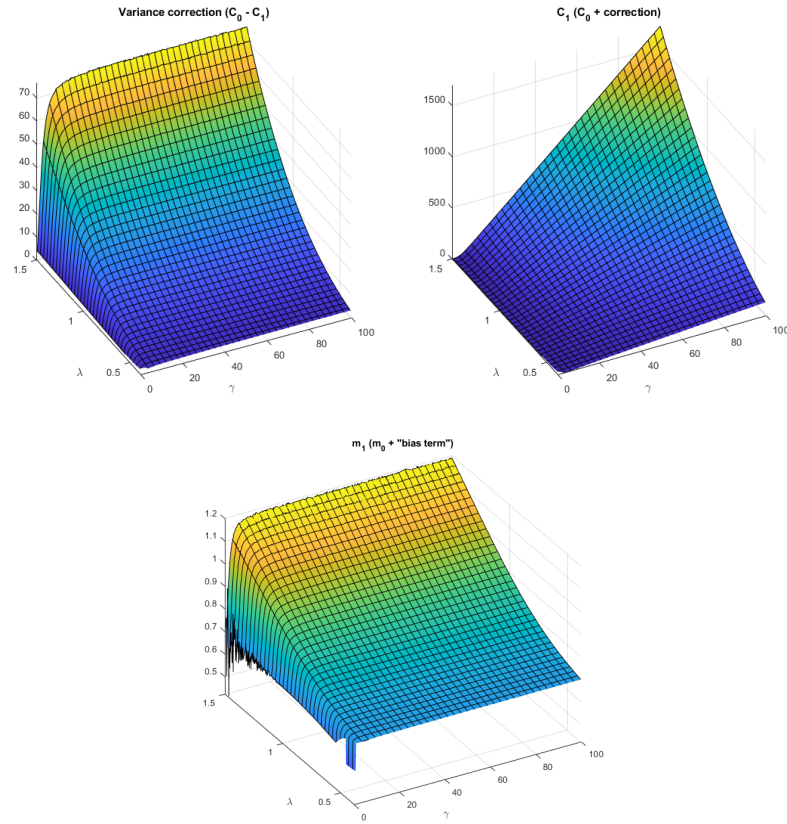


Figure 2: Behaviour of the diagnostic for a 1-dimensional T-distribution with 1 degree of freedom. Top left: variance correction term (difference between C_0 and C_1). Top right: posterior variance. Bottom: posterior mean.

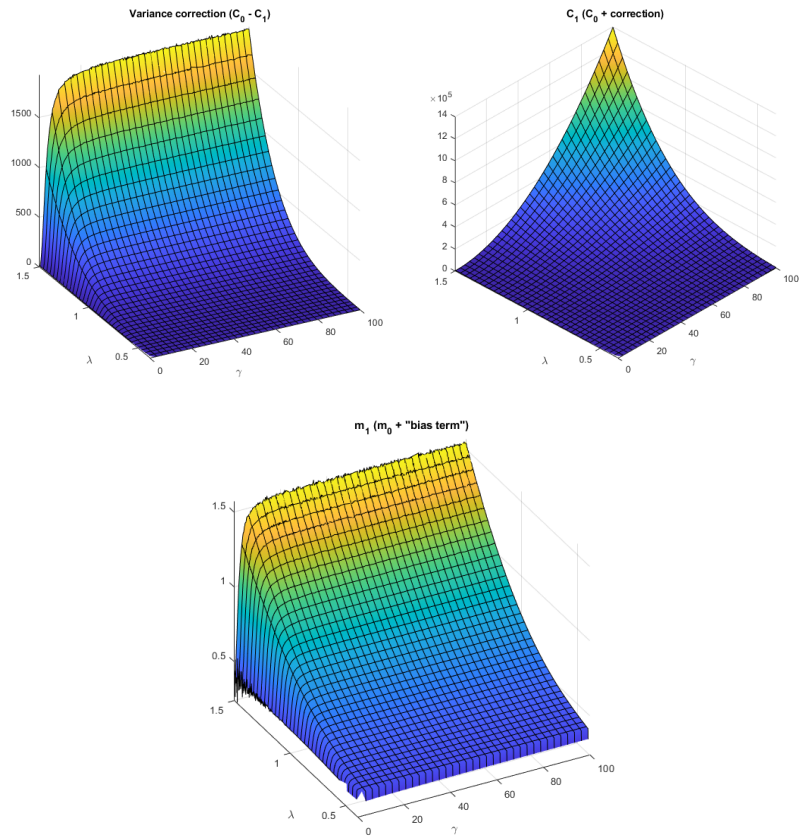


Figure 3: Behaviour of the diagnostic for a 2-dimensional T-distribution with 1 degree of freedom. Top left: variance correction term. Top right: posterior variance. Bottom: posterior mean.

approximation to be deemed reasonable. The second requirement ensures that the posterior GP provides a good fit to this threshold function. Although the diagnostic is not intended to produce high-accuracy approximations in general, at the very least it can be calibrated to do so for this special case. This is particularly important as the choice of hyperparameters can compromise inference in ways that require more in-depth scrutiny to understand, as will be discussed in the following subsection.

6.1 Calibrating in two dimensions

Recall that the diagnostic is intended to test whether or not the use of the Laplace approximation is justified by the shape of the true function. To this end, the actual values produced by the integral posterior do not necessarily tell the whole story. Note that the posterior integral mean m_1 (8) and GP mean m_1^t (5) are related by $m_1 = \int_{\mathbb{R}^d} m_1^t(t) dt$. Thus, the hyperparameters λ and γ can affect inference in ways not immediately obvious from Figures 2 and 3, by influencing the shape of the GP posterior mean.

With these considerations in mind, let us attempt to calibrate the diagnostic in two dimensions using the bivariate T distribution with $\nu_2 = 38$ degrees of freedom. The effects of λ and γ on m_1^t and the resulting inference for the integral will be assessed, with α set to put the Laplace approximation on the edge of the rejection region as described above. The interrogation grid will be expanded slightly from the one used for Figure 3, adding points at ± 3 standard deviations from the mode along each axis for a total of 13 interrogation points.

As in Figures 2-3, we found that γ had a diminishing effect on the shape of m_1^t as it increased beyond a certain threshold (not shown). Thus, we initially fix the moderately high (and somewhat arbitrary) value $\gamma = 0.25$ and vary λ , which controls the smoothness of the GP as it does with the conventional squared exponential kernel. For a low value of λ , Figure 4 shows the difference $m_1^t - T_{38,2}$ and the posterior inference for F . By (5), it holds that $m_1^t(\mathbf{s}) = f(\mathbf{s})$ for any f and any combination of hyperparameter values. However, at any other point t , the extent to which $m_1^t(t)$ updates from $m_0^t(t)$ is determined by the “weights” $C_t^0(t, \mathbf{s})^\top [C_t^0(\mathbf{s}, \mathbf{s})]^{-1}$. When λ is small, there is almost no prior dependence between GP values at distinct points, so these weights are close to zero for $t \notin \mathbf{s}$. This explains the spikes in Figure 4: the posterior mean is forced to equal f at the interrogation points, but everywhere else it is virtually unchanged from the prior mean. Thus, the posterior mean of the integral, m_1 , barely updates from the prior guess $L(T_{38,2})$. In this case, there is not much point in calibrating α as described above.

Figure 5 shows the effect of a moderate increase in λ . This induces more dependence between values at distinct points, and therefore more smoothness, in the GP. Thus, m_1^t updates to a greater extent along the lines defined by the interrogation points, and the value of m_1 moves closer to the true F . However, m_1^t still does not adequately update in regions outside of the interrogation grid, particularly around the main diagonal lines in the plane. To allow sufficient interpolation of the interrogated values in these regions, it is necessary to

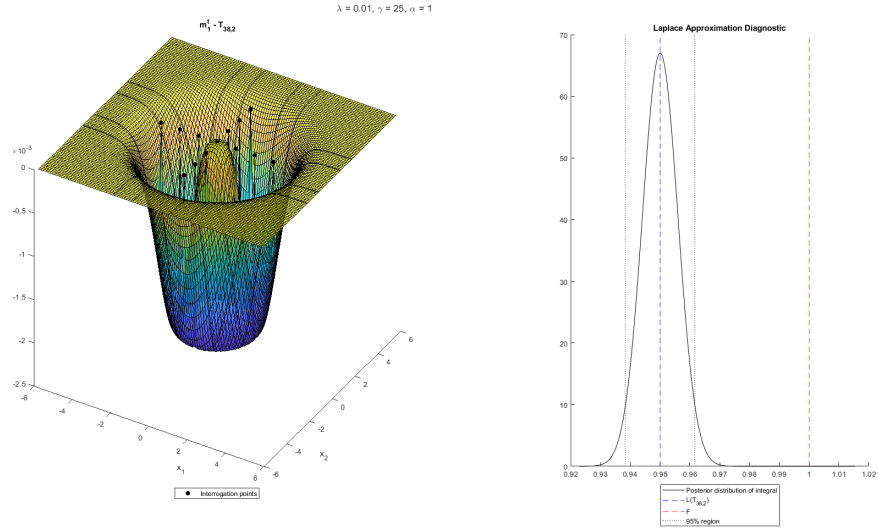


Figure 4: Posterior results in 2 dimensions with a high γ value and low λ value.

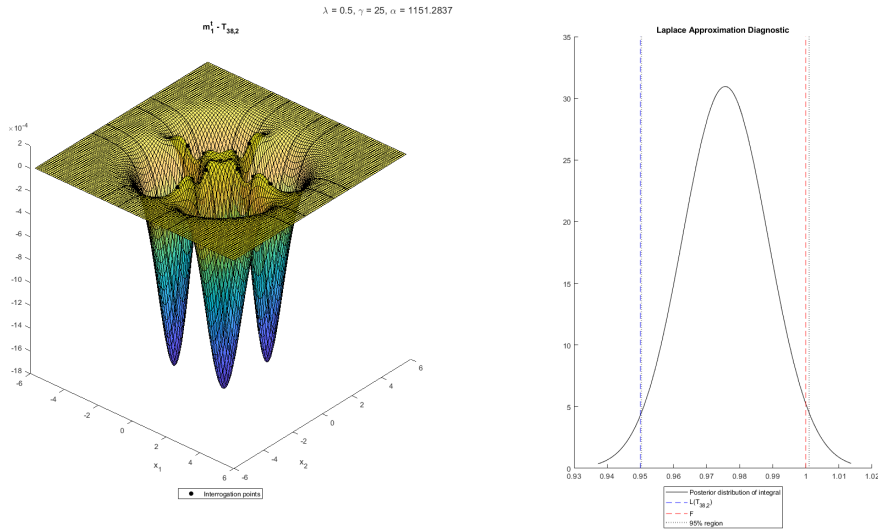


Figure 5: Posterior results in 2 dimensions with a high γ value and low λ value.

further increase λ as in Figure 6. However, this comes at a price: increasing λ to the necessary extent to update the posterior mean in the “diagonal regions” causes undesirable extrapolation effects due to oversmoothing. Indeed, in all four directions just beyond the extremal interrogation points, m_1^t dips well below $T_{38,2}$. As a result, m_1 is well *below* the Laplace approximation, let alone the true integral value. Oversmoothing causes the weights $C_t^0(t, \mathbf{s})^\top [C_t^0(\mathbf{s}, \mathbf{s})]^{-1}$ to have unpredictable effects beyond the boundaries of the interrogation grid, depending on the spread and density of \mathbf{s} as well as the shape of f . In some cases, the “valleys” seen in Figure 6 may be replaced by large “hills”, causing m_1 to significantly overestimate the value of F (not shown). In any case, it seems that good interpolation *within* the interrogation boundaries comes at the expense of poor extrapolation *beyond* them.

Of course, such extrapolation issues would not matter if we simply integrated only over the rectangle enclosed by the extremal interrogation points as in Zhou [2]. One may also wonder if the effects of oversmoothing could be mitigated by using a kernel which enforces less smoothness, such as the Matérn or uniform kernel [sources]. However, we have not yet exhausted all possible behaviours for the kernel (11), as the decay parameter γ can still be altered. Recalling that γ controls the covariance of points far away from the origin, it is reasonable to hypothesize that a lower γ -value would reduce extreme behaviour in the weights at points beyond the interrogation boundaries. In this case, good interpolation could then be achieved by increasing λ *without* sacrificing extrapolation. Indeed, attempting to minimize the L^2 -error $\int_{\mathbb{R}^d} (m_1^t(t) - T_{38,2}(t))^2 dt$ with respect to λ and γ (see Section 6.2) leads to a suggestion of $(\lambda, \gamma) \approx (5.3909, 0.8726)$. Figure 7 shows the results with these parameters. Some of the undesirable behaviour associated with both interpolation and extrapolation is present, but both types of behaviour are reduced in magnitude compared to Figures 5 and 6. As a result, the posterior inference for the integral is more accurate than it was with any other hyperparameter combination considered thus far.

6.2 On quantitative optimization of hyperparameters

6.3 The curse of dimensionality

Recall that we use a T -distribution with ν_d degrees of freedom to calibrate the d -dimensional diagnostic using the Laplace approximation formula (13). Although (13) increases towards 1 as $\nu \rightarrow \infty$ for fixed d , it is decreasing in d for fixed ν . Therefore, ν_d is an *increasing* function of d , as higher degrees of freedom are needed to ensure $L(\nu_d, d) \geq 0.95$ when d is large. Put another way, in higher dimensions, a T density must be closer in shape to a Gaussian for the Laplace approximation to be within 5% of the true integral.

This is a special case of a more general phenomenon: small differences in function shapes have more opportunity to influence their integrals in high dimensions. In heuristic terms, for large d differences between a function f and its Gaussian approximation m_0^t are compounded over a higher number of dimensions, thus increasing the discrepancy between their integrals. Therefore,

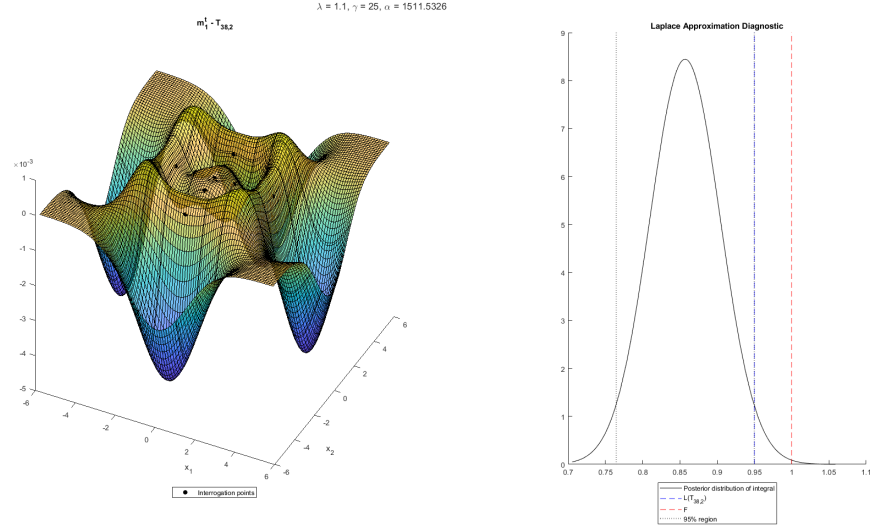


Figure 6: Posterior results in 2 dimensions with a high γ value and high λ value.

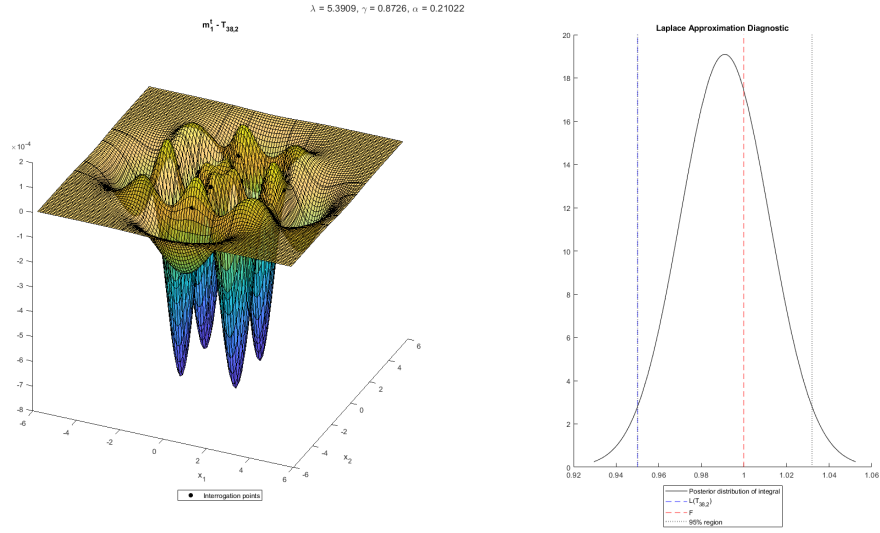


Figure 7: Posterior results in 2 dimensions with a low γ value and high λ value.

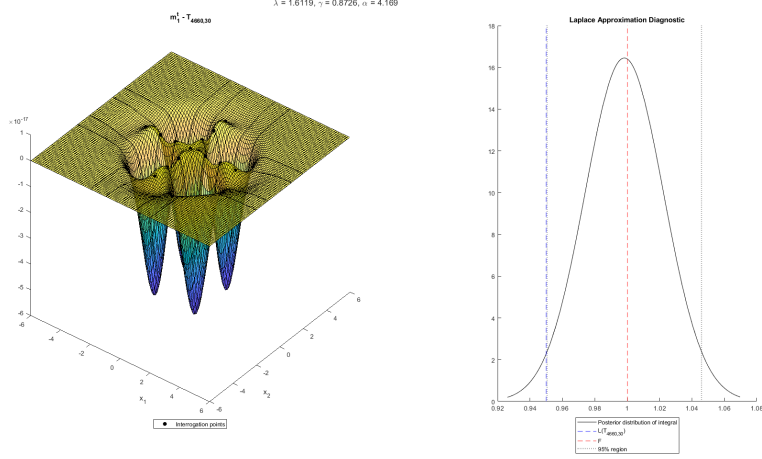


Figure 8: Posterior results for the empirically calibrated diagnostic in 30 dimensions. The left plot is a two-dimensional slice of $m_1^t - T_{4660,30}$, for which all 28 other coordinates were taken to be 0.

by necessity the diagnostic must be more shape-sensitive at high dimensions, as the Laplace approximation is reasonable for a diminished variety of function shapes. These high-dimensional considerations can result in some unusual diagnostic behaviour, as explained below.

Consider the diagnostic in 30 dimensions, calibrated with a T density with $\nu_{30} = 4660$ degrees of freedom. Rather than attempting another optimization based on L^2 -error (which involves a 30-dimensional numerical integral in this case), we simply used herusitics and visual inspection. Maintaining the γ value of 0.8726 used in two dimensions, λ was selected such m_1^t appeared as close as possible to $T_{4660,30}$, ultimately resulting in $(\lambda, \gamma, \alpha) = (1.6119, 0.8726, 4.169)$ (Figure 8).

Note that the multivariate T density used thus far corresponds to the joint distribution whose components are uncorrelated but *not* independent. The independent case results in a different density function: a product of d univariate T distributions,

$$f_{\nu,d}(t) = \prod_{i=1}^d \frac{\Gamma(\frac{\nu+1}{2})}{\Gamma(\frac{\nu}{2}) \sqrt{\nu\pi}} \left(1 + \frac{t_i}{\nu}\right)^{-\frac{\nu+1}{2}}. \quad (14)$$

It turns out that fewer degrees of freedom are required to bring the Laplace approximation of this function within 5% of the true integral (1): in particular, $L(f_{439,30}) = 0.95$. Figure 9 shows the surprising results of the diagnostic applied to $f_{439,30}$, using the empirically-determined hyperparameters stated above. The valleys typically seen in the “diagonal regions” are now large hills, causing m_1 to significantly overestimate F . This is seemingly unfortunate behaviour, given that the Laplace approximation is within 5% of the true integral and $f_{439,30}$ is

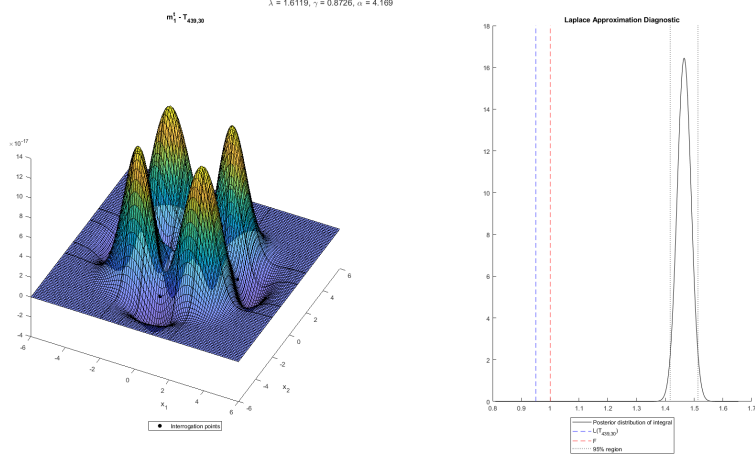


Figure 9: Posterior results for the empirically calibrated diagnostic in 30 dimensions. The left plot is a two-dimensional slice of $m_1^t - f_{439,30}$, for which all 28 other coordinates were taken to be 0.

visually indistinguishable from a Gaussian.

There are sensible reasons for this behaviour. Due to the placement of interrogation points, the diagnostic only uses shape information along the main axes of the domain. On these axes, $f_{439,30}$ differs from its Gaussian approximation to a greater proportional extent than $T_{4660,30}$. However, unlike the true multivariate T density, the product of independent densities does *not* have spherical contours, and in particular its values are lower at points which are not on the axes. Thus, the calculation of m_1^t uses the larger differences *on* the axes and overestimates the size of the correction needed in the diagonal regions. For a contrast, consider a different function

$$g_{\nu,d}(t) = \left[\frac{\Gamma(\frac{\nu+1}{2})}{\Gamma(\frac{\nu}{2}) \sqrt{\nu\pi}} \right]^d \left(1 + \frac{\|t\|}{\nu} \right)^{-\frac{\nu+1}{2}}. \quad (15)$$

This function is equal to $f_{\nu,d}$ on the axes, has the same Laplace approximation, and has spherical contours. However, the integral of $g_{439,30}$ over \mathbb{R}^{30} is equal to 1.661. The diagnostic does a good job of estimating *this* integral, but has no information encoding the differences between the two functions.

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