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Adaptive Design of Experiments for Conservative Estimation of Excursion Sets

Dario Azzimonti^{*†}, David Ginsbourger^{†*}, Clément Chevalier[‡]
Julien Bect[§], Yann Richet[¶]

Abstract

We consider a Gaussian process model trained on few evaluations of an expensive-to-evaluate deterministic function and we study the problem of estimating a fixed excursion set of this function. We review the concept of conservative estimates, recently introduced in this framework, and, in particular, we focus on estimates based on Vorob'ev quantiles. We present a method that sequentially selects new evaluations of the function in order to reduce the uncertainty on such estimates. The sequential strategies are first benchmarked on artificial test cases generated from Gaussian process realizations in two and five dimensions, and then applied to two reliability engineering test cases.

Keywords: Batch sequential strategies; Conservative estimates; Design of experiments; Excursion sets; Gaussian process models.

1 Introduction

The problem of estimating the set of inputs that leads a system to a particular behaviour is common in many applications, notably reliability engineering (see, e.g., Bect et al., 2012; Chevalier et al., 2014), climatology (see, e.g., French and Sain, 2013; Bolin and Lindgren, 2015) and many other fields (see, e.g., Bayarri et al., 2009; Arnaud et al., 2010; Wheeler et al., 2014). Here we consider a system modelled as a continuous function $f : \mathbb{X} \rightarrow \mathbb{Y}$, where \mathbb{X} is a locally compact Hausdorff topological space and $\mathbb{Y} = \mathbb{R}^k$, for some $k > 0$. Typical examples for \mathbb{X} are \mathbb{R}^d , differentiable manifolds or discrete spaces. In this work,

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we are interested in estimates for the set

$$\Gamma^* = \{x \in \mathbb{X} : f(x) \in T\}, \quad (1)$$

where $T \subset \mathbb{Y}$ is closed, which implies that Γ^* is closed as f is continuous. We focus on the case where few evaluations of the response f can be made available. This is common if f is expensive-to-evaluate, as, for example, when its evaluations require time consuming computer experiments (Sacks et al., 1989).

In a Bayesian framework (see, e.g., Chilès and Delfiner, 2012, and references therein) we assume that f is a realization of an almost surely continuous Gaussian process (GP) $Z \sim GP(\mathbf{m}, \mathfrak{K})$, with mean function \mathbf{m} , defined as $\mathbf{m}(x) := \mathbb{E}[Z_x]$, $x \in \mathbb{X}$, and covariance kernel $\mathfrak{K}(x, y)$, defined as $\mathfrak{K}(x, y) := \text{Cov}(Z_x, Z_y)$, $x, y \in \mathbb{X}$. For $n > 0$, we consider evaluations of f , $\mathbf{f}_{\mathbf{X}_n} = (f(x_1), \dots, f(x_n)) \in \mathbb{Y}^n$, at an initial design $\mathbf{X}_n = (x_1, \dots, x_n) \in \mathbb{X}^n$ and the posterior distribution of the process conditioned on the event $Z_{\mathbf{X}_n} = \mathbf{f}_{\mathbf{X}_n}$, where $Z_{\mathbf{X}_n} = (Z_{x_1}, \dots, Z_{x_n})$. We denote with $\mathbf{m}_n(x) = \mathbb{E}[Z_x \mid Z_{\mathbf{X}_n} = \mathbf{f}_{\mathbf{X}_n}]$, $x \in \mathbb{X}$, the posterior mean and with $\mathfrak{K}_n(x, x') = \text{Cov}(Z_x, Z_{x'} \mid Z_{\mathbf{X}_n} = \mathbf{f}_{\mathbf{X}_n})$, $x, x' \in \mathbb{X}$, the posterior covariance kernel. The prior distribution on Z induces a distribution on the excursion set

$$\Gamma = \{x \in \mathbb{X} : Z_x \in T\}. \quad (2)$$

Analogously, the posterior distribution on Z induces a posterior distribution for Γ and by summarizing this distribution we obtain an estimate for Γ^* . Random closed sets do not have a unique definition of expectation, however it is possible to provide different notions of expectation and variability for such objects (see, e.g., Molchanov, 2005). The Vorob'ev expectation (Vorob'ev, 1984) is an example recently introduced in the GP framework (see, e.g., Chevalier et al., 2013). Here we focus on conservative estimates for Γ introduced by French and Sain (2013) and Bolin and Lindgren (2015).

In what follows, we consider a Borel σ -finite measure μ defined on \mathbb{X} and denote with \mathfrak{C} a family of closed subsets in \mathbb{X} . A *conservative estimate at level α* for Γ^* is a set $\text{CE}_{\alpha, n}$ defined as

$$\text{CE}_{\alpha, n} \in \arg \max_{C \in \mathfrak{C}} \{\mu(C) : P_n(C \subset \Gamma) \geq \alpha\}, \quad (3)$$

where $P_n(\cdot) = P(\cdot \mid Z_{\mathbf{X}_n} = \mathbf{f}_{\mathbf{X}_n})$. This type of set estimate is particularly interesting in problems where Γ^* is a set of safe configurations for a system. In this case, by choosing a high α , we provide an estimate for Γ^* that with high posterior probability is included in Γ . The optimization procedure in equation (3) can be very challenging to solve, and it crucially depend on the choice of the family \mathfrak{C} . A common choice (see, e.g., French and Sain, 2013; Bolin and Lindgren, 2015; Azzimonti and Ginsbourger, 2016) is a parametric family of nested sets depending on a one dimensional parameter. Here we also rely on a specific one dimensional family and, in Section 2, we provide motivation for this choice.

The main contribution of this work is a framework to evaluate the uncertainty on conservative estimates for Γ^* and the development of strategies to

sequentially reduce such uncertainty by adding new function evaluations. If the input space \mathbb{X} is a compact subset of \mathbb{R}^d and the output space $\mathbb{Y} = \mathbb{R}$, sequential strategies have already been proposed for contour lines (Ranjan et al., 2008; Bichon et al., 2008) by adapting the expected improvement algorithm criterion (Jones et al., 1998). In the case of excursion sets, Stepwise Uncertainty Reduction (SUR) strategies based on the set’s measure were introduced by Vazquez and Bect (2009) and Bect et al. (2012). More recently a fast parallel implementation of these strategies have been proposed (Chevalier et al., 2014) and applied to the problem of identifying the set Γ^* . Here we extend this framework to conservative estimates.

In Section 1.1 we briefly recall the set estimates preliminary to this work and, in Section 1.2, the previously introduced uncertainty reduction techniques. In Section 2 we study the conservative estimates and we motivate our choice for the family \mathfrak{C} . In Section 3 we define the metrics used to quantify the uncertainty on conservative estimates. In Section 4, we detail the proposed sequential strategies and their implementation in real valued case ($\mathbb{Y} = \mathbb{R}$). In Section 5 we first benchmark the strategies with Gaussian process realizations showing how the conservative property affects the estimates. We then apply the proposed strategies on two industrial test cases: a coastal flood problem and a nuclear criticality safety problem.

1.1 The Vorob’ev approach to excursion set estimation

The posterior distribution of Γ provides estimates for Γ^* and different ways to quantify the uncertainty on these estimates. See, e.g. Chevalier et al. (2014); Bolin and Lindgren (2015); Azzimonti et al. (2016) for more details.

Let us now briefly recall the Vorob’ev approach (see, e.g., Vorob’ev, 1984; Molchanov, 2005; Chevalier et al., 2013). We define the *coverage probability function* of a random closed set Γ as

$$p_\Gamma(x) = P(x \in \Gamma), \quad x \in \mathbb{X}.$$

In our case we consider the posterior coverage function $p_{\Gamma,n}$, where we consider the posterior probability P_n . The coverage function defines the family of *Vorob’ev quantiles*

$$Q_{n,\rho} = \{x \in \mathbb{X} : p_{\Gamma,n}(x) \geq \rho\}, \quad (4)$$

with $\rho \in [0, 1]$. These sets are closed for each $\rho \in [0, 1]$ as the coverage function is upper semi-continuous (see Molchanov, 2005, Proposition 1.34).

The level ρ can be selected in different ways. A plug-in approach is to choose $\rho = 0.5$; this leads to the *Vorob’ev median*. Another important type of set estimate based on the Vorob’ev quantiles is the *Vorob’ev expectation* of Γ , where the level $\rho = \rho_V$ is chosen such that Q_ρ has the closest possible measure $\mu(Q_\rho)$ to the expected measure of Γ , $\mathbb{E}[\mu(\Gamma)]$. In Section 2 we show that Vorob’ev quantiles are a reasonable choice for the family \mathfrak{C} in the conservative estimates defined in equation (3).

Finally let us introduce the concept of expected distance in measure between two random sets. Consider two random closed sets $\Gamma_1, \Gamma_2 \subset \mathbb{X}$. The *expected distance in measure* between Γ_1 and Γ_2 with respect to the measure μ is

$$d_\mu(\Gamma_1, \Gamma_2) = \mathbb{E}[\mu(\Gamma_1 \Delta \Gamma_2)],$$

where $\Gamma_1 \Delta \Gamma_2 = (\Gamma_1 \setminus \Gamma_2) \cup (\Gamma_2 \setminus \Gamma_1)$. The expected distance in measure can be seen as the measure of variability associated with a Vorob'ev quantile and with the Vorob'ev expectation. In the Ph.D. thesis of Chevalier (2013), this notion of variability was introduced to adaptively reduce the uncertainty on Vorob'ev expectations for expensive-to-evaluate functions. Here, it is used in Section 3 to provide uncertainty functions for conservative estimates.

1.2 Background on SUR strategies

The objective of sequential strategies for GP models is to select a sequence of points X_1, X_2, \dots, X_n that reduces the uncertainty on selected posterior quantities. Here we are interested in reducing the uncertainty on $\text{CE}_{\alpha, n}$, as defined in equation (3). We consider $\mathbf{X}_n \in \mathbb{X}^n$ and $\mathbf{f}_{\mathbf{X}_n} \in \mathbb{Y}^n$, with \mathbf{X}_n the locations where the function f was evaluated and $\mathbf{f}_{\mathbf{X}_n}$ the actual evaluations. In the remainder of the paper we denote with $\mathbb{E}_n[\cdot] = \mathbb{E}[\cdot \mid Z_{\mathbf{X}_n} = \mathbf{f}_{\mathbf{X}_n}]$ the expectation conditioned on the event $Z_{\mathbf{X}_n} = \mathbf{f}_{\mathbf{X}_n}$. We focus on Stepwise Uncertainty Reduction (SUR) strategies. A SUR strategy (see, e.g., Fleuret and Geman, 1999; Bect et al., 2012; Chevalier et al., 2014) selects the next evaluation in order to reduce a particular uncertainty function.

Consider a model where n evaluations of f are given. An uncertainty function for a particular estimate is here defined as a map

$$H_n : (\mathbb{X} \times \mathbb{Y})^n \rightarrow \mathbb{R},$$

that associates to each vector of couples $(x_i, Z_{x_i})_{i=1, \dots, n}$ a real value representing the uncertainty associated with the selected estimate. Since they are selected sequentially, both the locations X_1, \dots, X_n and the evaluations are random. More specifically, denote with \mathcal{A}_n the σ -algebra generated by the couples $X_1, Z_{X_1}, \dots, X_n, Z_{X_n}$. We denote with \mathcal{H}_n the \mathcal{A}_n measurable random variable that returns the uncertainty associated with \mathcal{A}_n , the σ -algebra generated by $(X_i, Z_{X_i})_{i=1, \dots, n}$.

If we assume that the first n evaluations of the field are known, a SUR strategy selects the locations $X_{n+1}^*, \dots, X_{n+q}^*$ that minimize $\mathbb{E}_n[\mathcal{H}_{n+q}]$, the future uncertainty in expectation. For a more complete and theoretical perspective on SUR strategies see, e.g., Bect et al. (2016) and references therein. There are many ways to proceed with the minimization introduced above, see, e.g., Osborne et al. (2009); Ginsbourger and Le Riche (2010); Bect et al. (2012); González et al. (2016) and references therein. Here we focus on sub-optimal strategies, also called batch-sequential *one-step lookahead strategies*, that select the next batch of locations by greedily minimizing the expected uncertainty at the next step. This choice is often justified in practice because it is possible to run the evaluations of the function in parallel thus saving wall-clock time.

Definition 1 (batch sequential one-step lookahead criterion). *We call batch sequential one-step lookahead sampling criterion a function $J_n : (\mathbb{X})^q \rightarrow \mathbb{R}$ that associates to each batch of q points $\mathbf{x}^{(q)} := (x_{n+1}, \dots, x_{n+q}) \in \mathbb{X}^q$ the expected uncertainty at the next step assuming this batch is evaluated*

$$J_n(\mathbf{x}^{(q)}) = \mathbb{E}_n [\mathcal{H}_{n+q} \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q}].$$

In Sections 3 and 4, we revisit the concepts of uncertainty function and SUR criterion for the problem of computing conservative estimates of an excursion set. In the next section instead we briefly review some properties of the conservative estimates defined in equation 3.

2 Properties of conservative estimate

The conservative estimate introduced in equation (3) requires the specification of the family \mathfrak{C} where to search for the optimal set CE_α . In practice, it is convenient to choose a parametric family indexed by a parameter $\theta \in \mathbb{R}^k$. Consider a nested family \mathfrak{C}_θ indexed by a real number $\theta \in [0, 1]$, i.e. for each $\theta_1 > \theta_2$

$$C_{\theta_1} \subset C_{\theta_2}, \tag{5}$$

for any $C_{\theta_1}, C_{\theta_2} \in \mathfrak{C}_\theta$. Let us define $C_0 = \mathbb{X}$ and assume that $\mu(\mathbb{X}) < \infty$. This is often the case as \mathbb{X} is either chosen as a compact subset of \mathbb{R}^d with μ the Lebesgue measure or μ is a probability measure on \mathbb{X} .

For each θ , we can define the function $\varphi_\mu : [0, 1] \rightarrow [0, +\infty)$ that associates to each $\theta \in [0, 1]$ the value $\varphi_\mu(\theta) := \mu(C_\theta)$, with $C_\theta \in \mathfrak{C}_\theta$. It is a non increasing function of θ because the sets C_θ are nested. We further define the function $\psi_\Gamma : [0, 1] \rightarrow [0, 1]$ that associates to each θ the probability $\psi_\Gamma(\theta) := P(C_\theta \subset \Gamma)$. The function ψ_Γ is non decreasing in θ due to the nested property in equation (5). In this set-up the computation of CE_α is reduced to finding the smallest $\theta = \theta^*$ such that $\psi_\Gamma(\theta^*) \geq \alpha$, which can be achieved with a simple dichotomic search.

The Vorob'ev quantiles introduced in equation (4) are a family of closed sets that satisfy the property in equation (5). Moreover they have the important property of being the minimizers of the expected distance in measure among sets with the same measure.

Proposition 1. *Consider a measure μ such that $\mu(\mathbb{X}) < \infty$. The Vorob'ev quantile*

$$Q_\rho = \{x \in \mathbb{X} : p_\Gamma(x) \geq \rho\}$$

minimizes the expected distance in measure with Γ among measurable sets M such that $\mu(M) = \mu(Q_\rho)$, i.e.,

$$\mathbb{E}[\mu(Q_\rho \Delta \Gamma)] \leq \mathbb{E}[\mu(M \Delta \Gamma)], \tag{6}$$

for each measurable set M such that $\mu(M) = \mu(Q_\rho)$.

Proof. see Appendix A □

The Vorob'ev quantiles have the smallest expected distance in measure with respect to Γ among sets with the same measure. They are thus an optimal family for conservative estimates with respect to the expected distance in measure. In general, however, the Vorob'ev quantile chosen for CE_α with this procedure is not the set S with the largest measure satisfying the property $P(S \subset \Gamma) \geq \alpha$. See supplementary material for a counterexample.

In the remainder of the paper we always consider \mathfrak{C} as the family of Vorob'ev quantiles. Given an initial design \mathbf{X}_n we can exploit the previously described properties and obtain $\text{CE}_{\alpha,n}$, a conservative estimate at level α for Γ^* . This estimate will also be denoted with Q_{n,ρ_n^α} , where ρ_n^α is the conservative Vorob'ev threshold. In the next section we introduce different ways to quantify the uncertainty on this estimate, while in Section 4 we introduce sequential strategies to reduce this uncertainty by adding new evaluations to the model.

3 Uncertainty quantification on $\text{CE}_{\alpha,n}$

In this section we introduce several uncertainty functions for conservative estimates. Here we always assume that n evaluations of f are available.

Our object of interest is Γ^* , therefore we require uncertainty functions that take into account the whole set. Chevalier and co-authors (Chevalier et al., 2013; Chevalier, 2013) evaluate the uncertainty on the Vorob'ev expectation with the expected distance in measure between the current estimate Q_{n,ρ_n} and the set Γ . Let us recall here that the *Vorob'ev uncertainty* of the quantile Q_{ρ_n} is the quantity

$$\mathcal{H}_n(\rho_n) = \mathbb{E}_n[\mu(\Gamma \Delta Q_{n,\rho_n})]. \quad (7)$$

In the following sections, this uncertainty quantification function is applied to the Vorob'ev expectation, $\rho_n = \rho_{V,n}$, to the Vorob'ev median, $\rho_n = 0.5$, and to the conservative estimate at level α , $\rho_n = \rho_n^\alpha$. The thresholds $\rho_{V,n}$ and ρ_n^α are \mathcal{A}_n measurable random variables, as they depend on the locations and observations $(X_i, Z_{X_i})_{i=1,\dots,n}$.

Consider a threshold $\rho \in [0, 1]$, and note that, by definition, the symmetric difference can be written as

$$\mathbb{E}_n[\mu(\Gamma \Delta Q_{n,\rho})] = \mathbb{E}_n[\mu(Q_{n,\rho} \setminus \Gamma)] + \mathbb{E}_n[\mu(\Gamma \setminus Q_{n,\rho})]. \quad (8)$$

Let us denote with $G_n^{(1)}(\rho) = \mu(Q_{n,\rho} \setminus \Gamma)$ the random variable associated with the measure of the first set difference and with $G_n^{(2)}(\rho) = \mu(\Gamma \setminus Q_{n,\rho})$ the random variable associated with the second one.

Remark 1. Consider the conservative estimate Q_{n,ρ_n^α} , then the ratio between the error $\mathbb{E}_n[G_n^{(1)}(\rho_n^\alpha)]$ and the measure $\mu(Q_{n,\rho_n^\alpha})$ is bounded by $1 - \alpha$, the chosen level for the conservative estimates. See Appendix A for a proof.

A conservative estimate Q_{n,ρ_n^α} aims at controlling the error $\mathbb{E}_n[G_n^{(1)}(\rho_n^\alpha)]$. With a broad use of the hypothesis testing lexicon we denote *Type I* error

at state n the quantity $\mathbb{E}_n[G_n^{(1)}(\rho_n^\alpha)]$ and *Type II* error at state n the quantity $\mathbb{E}_n[G_n^{(2)}(\rho_n^\alpha)]$. Type II error defines the following uncertainty function for $\text{CE}_{\alpha,n}$.

Definition 2 (Type II uncertainty). *Consider the Vorob'ev quantile Q_{n,ρ_n^α} corresponding to the conservative estimate at level α for Γ . The Type II uncertainty is the uncertainty function $\mathcal{H}_n^{\text{T2}}$ defined as*

$$\mathcal{H}_n^{\text{T2}}(\rho_n^\alpha) := \mathbb{E}_n[G_n^{(2)}(\rho_n^\alpha)] = \mathbb{E}_n[\mu(\Gamma \setminus Q_{n,\rho_n^\alpha})]. \quad (9)$$

Conservative estimates at high levels α tend to select regions inside Γ , by definition. In particular if the number of function evaluations is high enough to have a good approximation of the function f , the conservative estimates with high α tend to be inside the true excursion set Γ^* . In these situations the expected type I error is usually very small, as shown in Remark 1, while type II error could be rather large. Type II uncertainty is thus a relevant quantity when evaluating conservative estimates. In the test case studies we also compute the expected type I error to check that it is consistently small.

The last uncertainty function introduced in this section is the expected difference between the measure of Γ and the measure of Q_{n,ρ_n^α} .

Definition 3 (Uncertainty MEAS). *We denote the uncertainty function related to the measure μ with $\mathcal{H}_n^{\text{MEAS}}$, defined as*

$$\mathcal{H}_n^{\text{MEAS}}(\rho_n^\alpha) := \mathbb{E}_n[\mu(\Gamma) - \mu(Q_{n,\rho_n^\alpha})] \quad (10)$$

This quantity is a reasonable uncertainty function only for conservative estimates. In this case, in fact, this quantity is equal to $\mathbb{E}_n[G_n^{(2)}(\rho_n^\alpha) - G_n^{(1)}(\rho_n^\alpha)]$ and, if the estimate is completely included in Γ , then it is the Type II uncertainty.

4 SUR strategies for conservative estimates

We now consider a current design of experiments \mathbf{X}_n , for some $n > 0$ and we introduce one-step lookahead SUR strategies for conservative estimates. These strategies select the next batch of $q > 0$ locations $X_{n+1}, \dots, X_{n+q} \in \mathbb{X}$ in order to reduce the future uncertainty \mathcal{H}_{n+q} defined in equation (7), (9) and (10). We denote with \mathcal{A}_{n+q} the σ -algebra generated by the sequence $(X_i, Z_{X_i})_{i=1, \dots, n+q} \in (\mathbb{X} \times \mathbb{Y})^{n+q}$. The uncertainty \mathcal{H}_{n+q} is a \mathcal{A}_{n+q} measurable random variable that depends on the unknown response at the points X_{n+1}, \dots, X_{n+q} . In this section we define the SUR criteria for a conservative level ρ_{n+q}^α , which is a \mathcal{A}_{n+q} measurable random variable because it also depends on the unknown responses. While the criteria are properly defined in this case, there are no closed form formulae for the level ρ_{n+q}^α . For this reason, in the next section, we implement the criteria with the last computed level ρ_n^α .

We consider three sampling criterion based on the uncertainty functions introduced in equation (7), (9) and (10). The first is an adaptation of the

Vorob'ev criterion introduced in Chevalier (2013) and based on the Vorob'ev deviation (Vorob'ev, 1984; Molchanov, 2005; Chevalier et al., 2013).

$$\begin{aligned} J_n(\mathbf{x}^{(q)}; \rho_{n+q}^\alpha) &= \mathbb{E}_n [\mathcal{H}_{n+q}(\rho_{n+q}^\alpha) \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q}] \quad (11) \\ &= \mathbb{E}_n \left[\mu(\Gamma \Delta Q_{n+q, \rho_{n+q}^\alpha}) \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q} \right] \end{aligned}$$

for $\mathbf{x}^{(q)} = (x_{n+1}, \dots, x_{n+q}) \in \mathbb{X}^q$, where $Q_{n+q, \rho_{n+q}^\alpha}$ is the Vorob'ev quantile obtained with $n+q$ evaluations of the function at level ρ_{n+q}^α , the conservative level obtained with $n+q$ evaluations.

In the case of conservative estimates with high level α , each term of equation (8) does not contribute equally to the expected distance in measure, as observed in Remark 1. It is thus reasonable to consider a criterion to reduce the Type II uncertainty introduced in Definition 2.

$$\begin{aligned} J_n^{\text{T}2}(\mathbf{x}^{(q)}; \rho_{n+q}^\alpha) &= \mathbb{E}_n [\mathcal{H}_{n+q}^{\text{T}2}(\rho_{n+q}^\alpha) \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q}] \quad (12) \\ &= \mathbb{E}_n \left[G_n^{(2)}(Q_{n+q, \rho_{n+q}^\alpha}) \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q} \right], \end{aligned}$$

for $\mathbf{x}^{(q)} \in \mathbb{X}^q$. In Section 4.1 this criterion is derived for a fixed, \mathcal{A}_n measurable level $\rho_n \in [0, 1]$, for the case $\mathbb{X} \subset \mathbb{R}^d$ and $\mathbb{Y} = \mathbb{R}$.

The last criterion studied here for conservative estimates relies on the uncertainty function $\mathcal{H}_n^{\text{MEAS}}$. We can define the measure based criterion as

$$J_n^{\text{MEAS}}(\mathbf{x}^{(q)}; \rho_{n+q}^\alpha) = \mathbb{E}_n \left[\mu(\Gamma) - \mu(Q_{n+q, \rho_{n+q}^\alpha}) \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q} \right]. \quad (13)$$

Since we are interested in minimizing this criterion and $\mathbb{E}_n[\mu(\Gamma)]$ is independent from $\mathbf{x}^{(q)}$, we consider the equivalent function to maximize

$$\widetilde{J}_n^{\text{MEAS}}(\mathbf{x}^{(q)}; \rho_{n+q}^\alpha) = \mathbb{E}_n \left[\mu(Q_{n+q, \rho_{n+q}^\alpha}) \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q} \right].$$

Note that this criterion select points that are meant to increase the measure of the estimate and it is only reasonable for conservative estimates where the conservative condition on $Q_{n+q, \rho_{n+q}^\alpha}$ leads to sets with finite measure in expectation.

4.1 Implementation

In this section we detail the algorithmic aspects of the criteria.

The notions of an estimate's uncertainty and of sequential criteria can be defined in the generic setting introduced in Section 1.2, however in order to provide practical implementations of the criteria we need to restrict ourselves to a more specific framework. Here we fix $\mathbb{X} \subset \mathbb{R}^d$, a compact subset of \mathbb{R}^d , and $\mathbb{Y} = \mathbb{R}$. These choices are common especially in engineering and other scientific applications. We assume that Z is a GP with constant prior mean \mathbf{m} and covariance kernel \mathfrak{K} . Finally we derive formulas for the criteria for the set

$\Gamma^* = \{x \in \mathbb{X} : f(x) \in T\}$ with $T = [t, +\infty)$, where $t \in \mathbb{R}$ is a fixed threshold. It is straightforward to compute the criteria for $T = (-\infty, t]$ and to extend them for unions of bounded intervals.

The formulas for the criteria introduced here all rely on the posterior coverage probability function p_n , where the subscript Γ is dropped as the set is clear from the context. In particular, from the assumptions previously introduced it follows that, for each $n \geq 0$, $p_n(x) = \Phi((\mathbf{m}_n(x) - t)/s_n(x))$, with $x \in \mathbb{X}$, where Φ is the univariate standard Normal distribution, \mathbf{m}_n is the posterior mean of the process and $s_n(x) = \sqrt{\mathfrak{K}_n(x, x)}$ for all $x \in \mathbb{X}$.

The first criterion introduced in Section 4 is based on the symmetric difference between the set Γ and the Vorob'ev quantile $Q_{n,\rho}$. In Chevalier (2013), Chapter 4.2, the formula for this criterion in this framework was derived for the Vorob'ev expectation, i.e. the quantile at level $\rho = \rho_{n,V}$. In the following remark we first extend this result to an \mathcal{A}_n measurable quantile ρ_n .

Remark 2 (Criterion J_n). Under the previously introduced assumptions the criterion J_n can be expanded in closed-form as

$$\begin{aligned} J_n(\mathbf{x}^{(q)}; \rho_n) &= \mathbb{E}_n [\mu(\Gamma \Delta Q_{n+q, \rho_n}) \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q}] \\ &= \int_{\mathbb{X}} \left(2\Phi_2 \left(\begin{pmatrix} a_{n+q}(u) \\ \Phi^{-1}(\rho_n) - a_{n+q}(u) \end{pmatrix}; \begin{pmatrix} 1 + \gamma_{n+q}(u) & -\gamma_{n+q}(u) \\ -\gamma_{n+q}(u) & \gamma_{n+q}(u) \end{pmatrix} \right) \right. \\ &\quad \left. - p_n(u) + \Phi \left(\frac{a_{n+q}(u) - \Phi^{-1}(\rho_n)}{\sqrt{\gamma_{n+q}(u)}} \right) \right) d\mu(u), \end{aligned} \quad (14)$$

where

$$\begin{aligned} a_{n+q}(u) &= \frac{\mathbf{m}_n(u) - t}{s_{n+q}(u)}, & \mathbf{b}_{n+q}(u) &= \frac{K_q^{-1} \mathfrak{K}_n(\mathbf{x}^{(q)}, u)}{s_{n+q}(u)}, \\ \gamma_{n+q}(u) &= \mathbf{b}_{n+q}^T(u) K_q \mathbf{b}_{n+q}(u), \quad u \in \mathbb{X} \end{aligned} \quad (15)$$

with $\mathfrak{K}_n(\mathbf{x}^{(q)}, u) = (\mathfrak{K}_n(x_{n+1}, u), \dots, \mathfrak{K}_n(x_{n+q}, u))^T$. K_q is covariance matrix with elements $[\mathfrak{K}_n(x_{n+i}, x_{n+j})]_{i,j=1,\dots,q}$, $\Phi_2(\cdot; \Sigma)$ is the bivariate centred Normal distribution with covariance matrix Σ .

The proof of the previous remark is a simple adaptation of the proof in Chevalier (2013), Chapter 4.2.

Proposition 2. *The criterion $J_n^{\text{T}2}(\cdot; \rho_n^\alpha)$ can be expanded in closed-form as*

$$\begin{aligned} J_n^{\text{T}2}(\mathbf{x}^{(q)}; \rho_n^\alpha) &= \mathbb{E}_n \left[G_n^{(2)}(Q_{n+q, \rho_n^\alpha}) \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q} \right] \\ &= \int_{\mathbb{X}} \Phi_2 \left(\begin{pmatrix} a_{n+q}(u) \\ \Phi^{-1}(\rho_n^\alpha) - a_{n+q}(u) \end{pmatrix}; \begin{pmatrix} 1 + \gamma_{n+q}(u) & -\gamma_{n+q}(u) \\ -\gamma_{n+q}(u) & \gamma_{n+q}(u) \end{pmatrix} \right) d\mu(u). \end{aligned} \quad (16)$$

Proof. The proof is a simple adaptation of the Remark 2. See Chevalier (2013). \square

Proposition 3. *The criterion J_n^{MEAS} can be expanded in closed-form as*

$$\begin{aligned} \widetilde{J}_n^{\text{MEAS}}(\mathbf{x}^{(q)}; \rho_n^\alpha) &= \mathbb{E}_n [\mu(Q_{n+q, \rho_n^\alpha}) \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q}] \\ &= \int_{\mathbb{X}} \Phi \left(\frac{a_{n+q}(u) - \Phi^{-1}(\rho_n^\alpha)}{\sqrt{\gamma_{n+q}(u)}} \right) d\mu(u). \end{aligned} \quad (17)$$

Proof. First of all notice that, for each $x \in \mathbb{X}$, the coverage function $p_{n+q}(\mathbf{x}^{(q)})$ can be written as

$$p_{n+q}(\mathbf{x}^{(q)})(x) = \Phi(a_{n+q}(x) + \mathbf{b}_{n+q}^T Y_q), \quad (18)$$

where a_{n+q} , \mathbf{b}_{n+q} are defined in equation (15) and $Y_q \sim N_q(0, K_q)$ is a q dimensional normal random vector. The indicator function of the set Q_{n+q, ρ_n^α} can be written as $\mathbb{1}_{p_{n+q}(x) \geq \rho_n^\alpha}$. By Tonelli's theorem we exchange the expectation with the integral over \mathbb{X} and we obtain

$$\begin{aligned} &\mathbb{E}_n \left[\mathbb{E} [\mu(Q_{n+q, \rho_n^\alpha}) \mid X_{n+1} = x_{n+1}, \dots, X_{n+q} = x_{n+q}] \right] \\ &= \int_{\mathbb{X}} \mathbb{E}_n [\mathbb{1}_{p_{n+q}(u) \geq \rho_n^\alpha}] d\mu(u) = \int_{\mathbb{X}} P_n(p_{n+q}(u) \geq \rho_n^\alpha) d\mu(u). \end{aligned}$$

By substituting the expression in equation (18) we obtain

$$\begin{aligned} \int_{\mathbb{X}} P_n(p_{n+q}(u) \geq \rho_n^\alpha) d\mu(u) &= \int_{\mathbb{X}} P_n(a_{n+q}(u) + \mathbf{b}_{n+q}^T Y_q \geq \Phi^{-1}(\rho_n^\alpha)) d\mu(u) \\ &= \int_{\mathbb{X}} \Phi \left(\frac{a_{n+q}(u) - \Phi^{-1}(\rho_n^\alpha)}{\sqrt{\gamma_{n+q}(u)}} \right) d\mu(u) \end{aligned}$$

□

For the practical implementation of the sampling criteria we exploit the kriging update formulas (Chevalier et al., 2014; Emery, 2009) for faster updates of the posterior mean and covariance when new evaluations are added.

The sampling criteria, implemented above in equation (14), (16) and (17), are used to select the next evaluations of the function f as a part of a larger algorithm that provides conservative estimates for Γ^* . See supplementary material for more details. The conservative level ρ_n^α is computed with the algorithm detailed by Azzimonti and Ginsbourger (2016).

5 Test cases

In this section we apply the proposed sequential uncertainty reduction methods to different test cases. First we develop a benchmark study with Gaussian process realizations to study the different behaviour of the proposed strategies. Then, we apply the methods to two reliability engineering test cases. In the first

Table 1: Strategies implemented in the test cases.

Strategy number	criterion	parameters
Benchmark 1	IMSE	
Benchmark 2	tIMSE	target= t
A	$J_n(\cdot; \rho_n)$	$\rho_n = 0.5$
B	$J_n(\cdot; \rho_n)$	$\rho_n = \rho_n^\alpha, \alpha = 0.95$
C	$\widetilde{J}_n^{\text{MEAS}}(\cdot; \rho_n)$	$\rho_n = \rho_n^\alpha, \alpha = 0.95$
D	$J_n^{\text{T}^2}(\cdot; \rho_n)$	$\rho_n = 0.5$
E	$J_n^{\text{T}^2}(\cdot; \rho_n^\alpha)$	$\alpha = 0.95$
F (hybrid strategy)	$J_n^{\text{T}^2}(\cdot; \rho_n^\alpha) + \text{IMSE}$	2 iterations IMSE, 1 iteration with E

test case, the set of interest represents the offshore conditions that do not lead the water level at the coast to be larger than a critical threshold above which flood would occur. In the second test case the set of interest is the set of safe parameters for nuclear storage facility.

In each test case we choose a tensor product covariance kernel from the Matérn family, see Rasmussen and Williams (2006, Chapter 4), for details on the parametrization.

All computations are carried out in the **R** programming language (R Core Team, 2016), with the packages `DiceKriging` (Roustant et al., 2012) and `DiceDesign` (Franco et al., 2013) for Gaussian modelling, `KrigInv` (Chevalier et al., 2014) for already existing sampling criterion and `ConservativeEstimates` (Azzimonti and Ginsbourger, 2016) to compute the conservative estimates.

5.1 Benchmark study: Gaussian processes

Let us first benchmark the different strategies introduced in Section 4 on Gaussian process realizations in two and five dimensions. The following setup is shared between the two test case. We consider the unit hypercube $\mathbb{X} = [0, 1]^d$, $d = 2, 5$ and we choose a Gaussian process $(Z_x)_{x \in \mathbb{X}} \sim GP(\mathbf{m}, \mathfrak{K})$. The GP has a constant prior mean $\mathbf{m} = 0$ and a tensor product Matérn covariance kernel with parameters fixed as detailed in Table 2. The objective is to obtain a conservative estimate at level $\alpha = 0.95$ for $\Gamma = \{x \in \mathbb{X} : Z_x \geq 1\}$. The measure of reference μ is the Lebesgue measure on \mathbb{X} . Here we test the strategies detailed in Table 1.

We consider an initial design of experiments $\mathbf{X}_{n_{\text{init}}}$, obtained with the function `optimumLHS` from the package `lhs` and we simulate the field at $\mathbf{X}_{n_{\text{init}}}$. The size n_{init} is chosen small to highlight the differences between the sequential strategies. We select the next evaluations by minimizing each sampling criteria detailed in Table 1. Each criterion is run for $n = 30$ iterations, updating the model with $q = 1$ new evaluations at each step. We consider m_{doe} different

Table 2: Test cases parameter choices.

Test case	d	covariance parameters	m_{doe}	n_{init}
GP	2	$\nu = 3/2, \theta = [0.2, 0.2]^T, \sigma^2 = 1$	10	3
GP	5	$\nu = 3/2, \theta = \sqrt{\frac{5}{2}}[0.2, 0.2, 0.2, 0.2, 0.2]^T, \sigma^2 = 1$	10	6
Costal	2	$\nu = 5/2$, MLE for $\theta, \sigma^2, \sigma_{\text{noise}}^2$	10	10
Nuclear	2	$\nu = 5/2$, MLE for $\theta, \sigma^2, \sigma_{\text{noise}}^2$	10	10

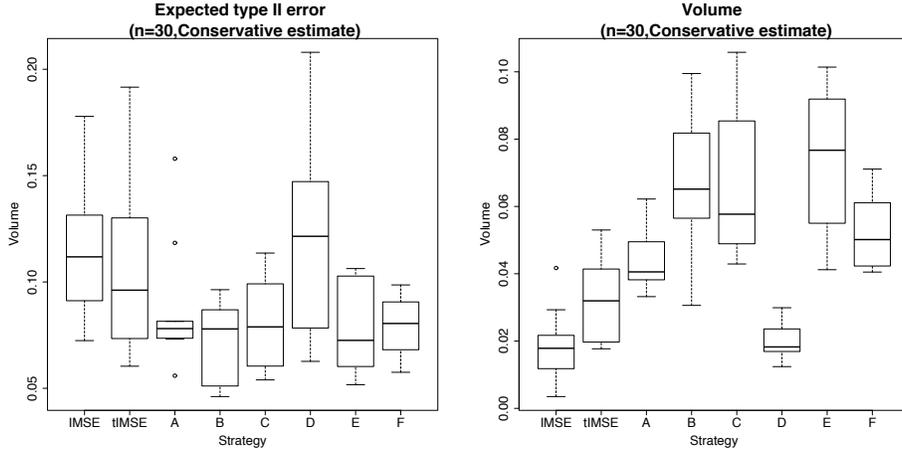


Figure 1: Median type II error for $Q_{\rho_n^\alpha}$ across the different designs of experiments, after $n = 30$ iterations. Test case in dimension 2. Figure 2: Measure $\mu(Q_{\rho_n^\alpha})$ across the different designs of experiments, after $n = 30$ iterations. Test case in dimension 2.

initial design of experiments and, for each design, we replicate the procedure 10 times with different initial values $Z_{\mathbf{X}_{n_{\text{init}}}}$.

5.1.1 Dimension 2

We evaluate the strategies by looking at the type I and type II errors for $Q_{\rho_n^\alpha}$, defined in Section 3, and by computing the measure $\mu(Q_{\rho_n^\alpha})$. For each of these quantities we report the median result over the replications obtained after $n = 30$ evaluations for each initial design.

Expected type I error does not vary much among the different strategies as it is controlled by the probabilistic condition imposed on the estimate, as shown in Section 2. See supplementary material for the results on expected type I error and the total computing time.

We show the distribution of expected type II error, in Figure 1, and the expected volume $\mu(Q_{\rho_n^\alpha})$ after $n = 30$ new evaluations, in Figure 2. The strategies A, B, C, E, F all provide better uncertainty reduction for conservative estimates

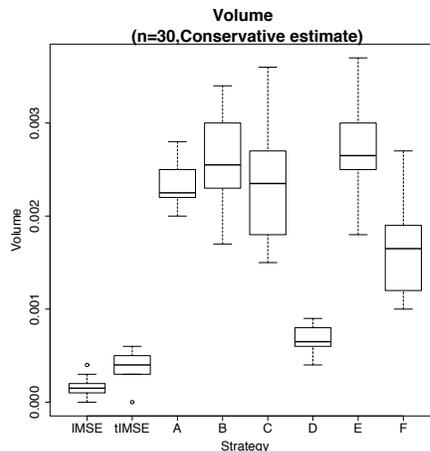
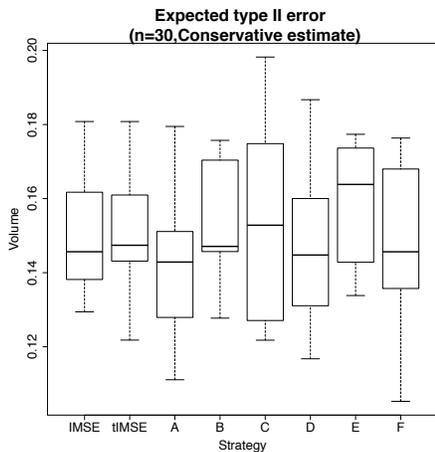


Figure 3: Median type II error for $Q_{\rho_n^\alpha}$ across different designs of experiments, after $n = 30$ iterations. Test case in dimension 5.

Figure 4: Measure $\mu(Q_{\rho_n^\alpha})$ across different designs of experiments, after $n = 30$ iterations. Test case in dimension 5.

than a standard IMSE strategy or than a tIMSE strategy. In particular strategy E has the lowest median type 2 error while at the same time providing an estimate with the largest measure, thus yielding a conservative set which is likely to be included in Γ^* and not too small in volume. All estimates are very conservative: the median ratio between the expected type I error and the estimate's volume is 0.03%, thus much smaller than the upper bound $1 - \alpha = 5\%$ computed in Remark 1. On the other hand the expected type II error is in median 178% bigger than the estimate volume.

5.1.2 Dimension 5

In Figures 3 and 4 we show the distribution of expected type II errors for $Q_{\rho_n^\alpha}$ and its measure $\mu(Q_{\rho_n^\alpha})$ obtained with the different design of experiments, after 30 iterations of each strategy. The resulting expected type I error and the total computational time are reported in the supplementary material.

In this test case the strategies are harder to rank. The IMSE strategy provides conservative estimates with small measure and with slightly larger type II error. Strategies A, B, C, E provide a good trade off between small type II error and large measure of the estimate, however they are not clearly better than the other strategies in this case. The estimates provided by all methods are very conservative also in this case. The median ratio over all DoEs and all replications between the expected type I error and volume is 0.33%, which is smaller than the upper bound 5%, as computed in Remark 1. The expected type II error is instead 3 orders of magnitude larger than the estimate's volume. This indicates that we have only recovered a small portion of the true set Γ^* ,

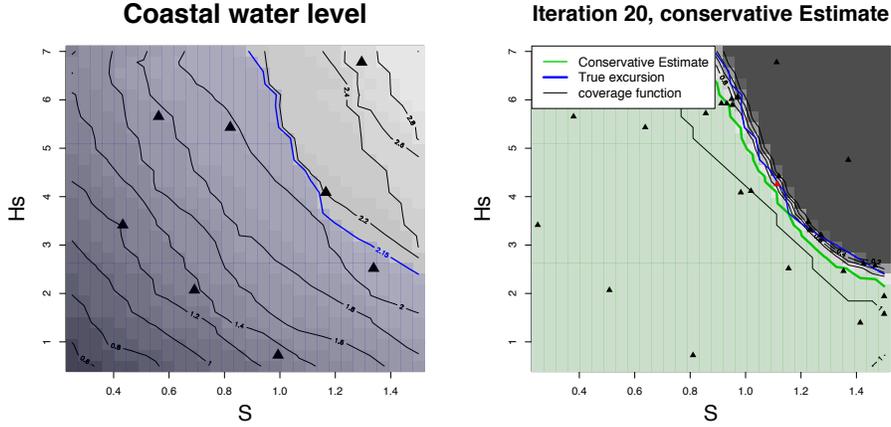


Figure 5: Coastal flood test case. Water level, set of interest (shaded blue, coverage function, conservative estimate volume=77.56%) and one initial design (shaded green, $\alpha = 0.95$) after 30 function evaluations (strategy E).

however, under the model, this estimate is very conservative.

5.2 Coastal flood test case

In this section we present a coastal flood test case introduced in Rohmer and Idier (2012). We focus here on studying the parameters that lead to floods on the coastlines. While similar studies are often conducted with full grid simulations, those techniques require many hours of computational time and, thus, render set estimation problems often infeasible. The use of meta-models, recently revisited in this field (see, e.g., Rohmer and Idier, 2012, and references therein), allows tackling this computational issue.

Here we consider a simplified coastal flood case as described by Rohmer and Idier (2012). The water level at the coast is modelled as a deterministic function $f : \mathbb{X} \subset \mathbb{R}^d \rightarrow \mathbb{R}$, assuming steady offshore conditions, without solving the flood itself inland. The input space $\mathbb{X} = [0.25, 1.50] \times [0.5, 7]$ are the variables storm surge magnitude S and significant wave height H_s . We are interested in recovering the set $\Gamma^* = \{x \in \mathbb{X} : f(x) \leq t\}$, with $t = 2.15$. In order to evaluate the quality of the meta-model, we rely on a grid experiment of 30×30 runs carried out by Rohmer and Idier (2012).

Here we consider a Gaussian process prior $(Z_x)_{x \in \mathbb{X}} \sim GP(\mathbf{m}, \mathfrak{K})$, with constant prior mean function and Matérn covariance kernel with $\nu = 5/2$. We assume that the function evaluations are noisy with zero noise mean and variance σ_{noise}^2 . We select $m_{\text{doe}} = 10$ different initial DoEs, with equal size $n_{\text{init}} = 10$. The initial designs are chosen with a maximin LHS design $\mathbf{X}_{n_{\text{init}}} = \{x_1, \dots, x_{n_{\text{init}}}\} \subset \mathbb{X}$ with the function `optimumLHS` from the package `lhs`. The covariance kernel

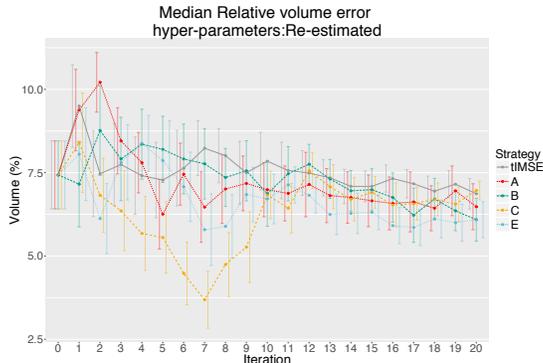
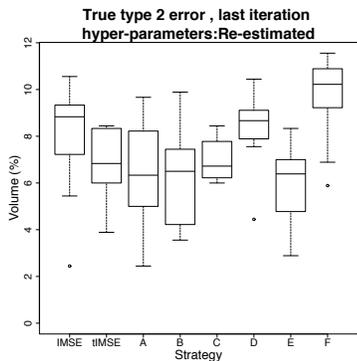


Figure 7: Coastal flood test case, randomized initial DoEs. True type II error for $CE_{\alpha,30}(\alpha = 0.95)$ at the last iteration. Figure 8: Coastal flood test case, randomized initial DoEs. Relative volume error as a function of iteration number. Strategies tIMSE, A, B, E.

hyper-parameters and the noise variance are estimated with maximum likelihood. Figure 5 shows the true function f we are aiming at reconstructing, the critical level and one initial design of experiments. We compute conservative set estimates for Γ^* at level $\alpha = 0.95$, as defined in Section 2, with the Lebesgue measure on \mathbb{X} .

We proceed to add 20 evaluations with the strategies detailed in Table 1. The covariance hyper-parameters are re-estimated at each step with maximum likelihood. See supplementary material for a study on the effect of parameter re-estimation. Figure 6 shows the conservative estimate obtained after 30 functions evaluations at locations chosen with Strategy E .

Figure 7 shows the true type II error at the last iteration of each strategy, after 30 evaluations of the function. The true type II error is computed by comparing the conservative estimate with an estimate of Γ^* obtained from the 30×30 grid experiment. Monte Carlo integration over this grid of evaluations leads to a volume of Γ^* equal to 77.56%.

At the last iteration, strategies A, B, E provide estimates with higher volume and lower type II error in median than IMSE and tIMSE. For example, the median type II error for Strategy E is 38% smaller than the IMSE type II error. For all strategies the true type I error is zero for almost all initial DoEs, thus indicating that all strategies lead to conservative estimates.

Figure 8 shows the behaviour of relative volume error as a function of the iteration number for Strategies tIMSE, A, B, E. The hyper-parameter re-estimation causes the model to be overconfident at the initial iterations, thus increasing the relative volume error. As the number of evaluations increases the hyper-parameter estimates become more stable and the relative error decreases as conservative estimates are better included in the true set.

5.3 Nuclear criticality safety test case

In this section we test the proposed strategies in a reliability engineering test case from the French Institute of nuclear safety (IRSN). The problem at hand concerns a nuclear storage facility and we are interested in estimating the set of parameters that lead to a safe storage of the material. This is closely linked to the production of neutrons. In fact, since neutrons are both the product and the initiator of nuclear reactions, an overproduction could lead to a chain reaction. The safety of a system is usually evaluated with the neutron multiplication factor, here called k -effective or k -eff : $\mathbb{X} \rightarrow [0, 1]$ defined on $\mathbb{X} = [0.2, 5.2] \times [0, 5]$. The two parameters represent the fissile material density, PuO_2 , and the water thickness, H_2O . We are interested in recovering the set of safe configurations

$$\Gamma^* = \{(\text{PuO}_2, \text{H}_2\text{O}) \in \mathbb{X} : k\text{-eff}(\text{PuO}_2, \text{H}_2\text{O}) \leq 0.92\},$$

where the threshold was chosen at 0.92 for safety reasons.

In general, the evaluation of k -eff at one point requires a MCMC simulation whose results have an heterogeneous noise variance. This is an expensive computer experiment thus our objective is to provide an estimate for Γ^* from few evaluations of k -eff and to quantify its uncertainty. The criteria implemented in Section 4.1 are not adapted to heterogeneous noise variance and could choose suboptimal locations. We could consider the noise homogeneous and estimate it from the data. However this procedure might lead to large errors in the estimates if the true noise variance is highly heterogeneous. In order to avoid such pitfalls we consider a smoothed version of k -eff. An approximation of k -eff is first computed from a 50×50 grid of evaluations of k -eff in \mathbb{X} . We then consider k -eff as a realization of a Gaussian process with mean zero, tensor product Matérn ($\nu = 5/2$) covariance kernel and heterogeneous noise variance equal to the MCMC variance and we compute the posterior mean of this field given the 2500 observations. We use this function as the true function. In what follows we denote with k -eff the result of this smoothing operation. Figure 9 shows this function and the set Γ^* .

Given the previous assumptions on k -eff we fix a prior $(Z)_{x \in \mathbb{X}} \sim GP(\mathbf{m}, \mathfrak{K})$ with constant mean function \mathbf{m} and tensor product Matérn covariance kernel with $\nu = 5/2$. Even if we consider the smoothed function, we still assume that evaluations are noisy with zero mean noise and variance σ_{noise}^2 for numerical stability. We consider $m_{\text{doe}} = 10$ different initial DoEs of size $n_0 = 10$, obtained with the function `optimumLHS` from the package `lhs` in **R**. Figure 9 one initial design of experiments. Each initial DoE is used to estimate the covariance hyper-parameters and the noise variance with maximum likelihood.

We now test how to adaptively reduce the uncertainty on the estimate with different strategies. Table 1 lists the strategies tested in this section. We run $n = 20$ iteration of each strategy and at each step we select a batch of $q = 3$ new points where to evaluate k -eff. The covariance hyper-parameters are re-estimated at each iteration by adding 3 new evaluations. The starting model provides a conservative estimate for Γ^* at level $\alpha = 0.95$, with the Lebesgue measure μ on \mathbb{X} . Figure 10 shows the coverage function of the random set Γ

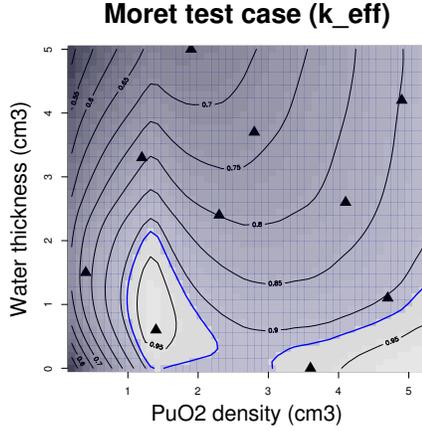


Figure 9: Nuclear criticality safety test case. Smoothed function k_{eff} , set of interest Γ^* (shaded blue, volume=88.16%) and one initial design of experiments.

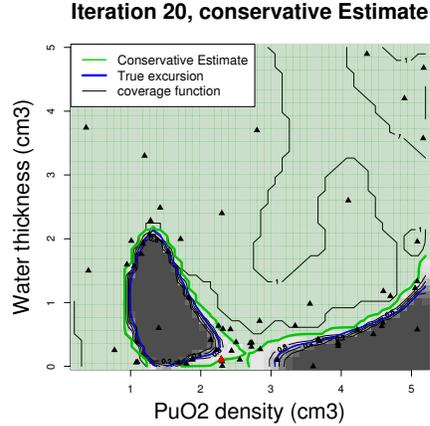


Figure 10: Nuclear criticality safety test case. Coverage function, conservative estimate ($\alpha = 0.95$, shaded green) after 70 function evaluations (strategy E).

obtained after 70 function evaluations at locations selected with Strategy E and the corresponding conservative estimate.

Figure 11 shows a comparison of the type II error at the last iteration, i.e. after 70 evaluations of the function, for each initial DoE and each strategy. Strategy D as in the previous test cases is the worse performer, thus showing that minimization of type II error works only for conservative quantiles. Strategies A, B, E perform well both in terms of final volume and true type II error. Strategy E , for example achieves a type II error 27% lower than IMSE. While still performing better than strategy D , F does not show any improvement over other strategies. Strategy C also performs well in this test case showing a 16% lower median type II error than IMSE.

Figure 12 shows the relative volume error as a function of the iteration number for strategies tIMSE, A, B, C, E . The relative volume error is computed by comparing the conservative estimate with an estimate of Γ^* obtained from evaluations of k_{eff} on a grid 50×50 . The volume of Γ^* computed with Monte Carlo integration from this grid of evaluations is 88.16%. All strategies presented show a strong decrease in relative volume error in the first 10 iteration, i.e. until 40 evaluations of k_{eff} are added. In particular strategies B, C, E show the strongest decline in error in the first 5 iterations. Overall, as in the previous test cases, strategy E , the minimization of the expected type II error, seem to provide the best uncertainty reductions both in terms of relative volume error and in terms of type II error.

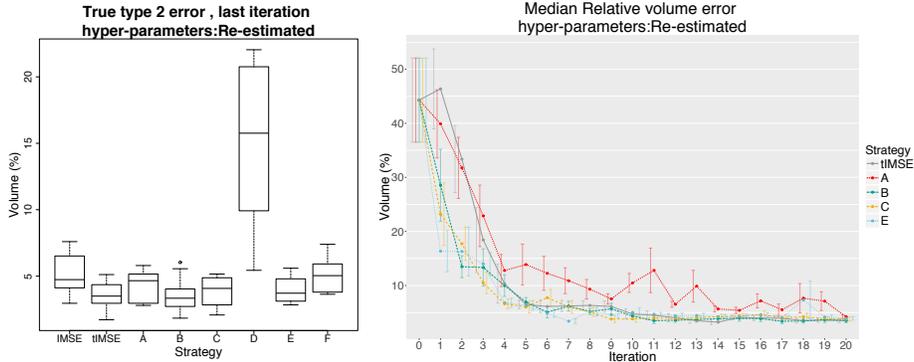


Figure 11: Nuclear criticality safety test case, randomized initial DoEs. True type II error at the final iteration. Figure 12: Nuclear criticality safety test case, randomized initial DoEs. Relative volume error as a function of the iteration number. Strategies tIMSE, A, B, C, E.

6 Discussion

In this paper we introduced sequential uncertainty reduction strategies for conservative estimates. This type of set estimates proved to be useful in reliability engineering, however they could be of interest in all situations where practitioners aim at controlling the overestimation of the set. The estimator CE, however, is based on a global quantity and an underlying GP model that badly approximates f will not lead to reliable estimates. For a fixed model, this issue might be reduced by increasing the level of confidence. We presented test cases with fixed $\alpha = 0.95$, however testing different levels, e.g. $\alpha = 0.99, 0.995$, and comparing the results is a good practice.

The sequential strategies proposed here provide a way to reduce the uncertainty on conservative estimates by adding new function evaluations. The numerical studies presented showed that adapted strategies provide a better uncertainty reduction than generic strategies. In particular, strategy E, i.e. the criterion $J_n^{T^2}(\cdot; \rho_n^\alpha)$, resulted among the best criteria in terms of Type 2 uncertainty and relative volume error in all test cases. In this work we mainly focused on showing the differences between the strategy with a-posteriori measures of uncertainty. Nonetheless the expected type I and II errors could be used to provide stopping criteria for the sequential strategies. Further studies in this direction are needed to understand the limit behaviour of these quantities as the number of evaluations increases.

The strategies proposed in this work focus on reducing the uncertainty on conservative estimates. This objective does not necessarily lead to better overall models for the function or to good covariance hyper-parameters estimation. The sequential behaviour of hyper-parameters maximum likelihood estimators under SUR strategies needs to be studied in more details. See supplementary material for a preliminary study on this aspect.

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A Properties of conservative estimates

In the following, let us denote with (Ω, \mathcal{F}, P) a probability space.

Proof of Proposition 1. Let us consider a measurable set M such that $\mu(M) = \mu(Q_\rho)$. For each $\omega \in \Omega$, we have

$$\begin{aligned} \mu(M\Delta\Gamma(\omega)) - \mu(Q_\rho\Delta\Gamma(\omega)) &= 2\left(\mu(\Gamma(\omega) \cap (Q_\rho \setminus M)) - \mu(\Gamma(\omega) \cap (M \setminus Q_\rho))\right) \\ &\quad + \mu(Q_\rho^C) - \mu(M^C). \end{aligned}$$

By applying the expectation on both sides and by remembering that $\mu(Q_\rho^C) = \mu(M^C)$ we obtain

$$\begin{aligned} \mathbb{E}[\mu(M\Delta\Gamma) - \mu(Q_\rho\Delta\Gamma)] &= \mathbb{E}\left[2\left(\mu(\Gamma \cap (Q_\rho \setminus M)) - \mu(\Gamma \cap (M \setminus Q_\rho))\right)\right] \\ &= 2 \int_{Q_\rho \setminus M} p_\Gamma(u) d\mu(u) - 2 \int_{M \setminus Q_\rho} p_\Gamma(u) d\mu(u), \end{aligned}$$

where the second equality comes from the definition of Q_ρ . Moreover, since $p_\Gamma(x) \geq \rho$ for $x \in Q_\rho \setminus M$ and $p_\Gamma(x) \leq \rho$ for $x \in M \setminus Q_\rho$ we have

$$\begin{aligned} 2 \left[\int_{Q_\rho \setminus M} p_\Gamma(u) d\mu(u) - \int_{M \setminus Q_\rho} p_\Gamma(u) d\mu(u) \right] &\geq 2\rho[\mu(Q_\rho \setminus M) - \mu(M \setminus Q_\rho)] \\ &= 2\rho[\mu(Q_\rho) - \mu(M)] = 0, \end{aligned}$$

which shows that Q_ρ verifies equation (6). \square

Proof of Remark 1. Notice that for all $\omega \in \Omega$ such that $Q_{n,\rho_n^\alpha} \subset \Gamma(\omega)$, we have $G_n^{(1)}(\omega) = 0$. By applying the law of total expectation we obtain

$$\begin{aligned} \mathbb{E}_n[G_n^{(1)}] &= \mathbb{E}_n[G_n^{(1)} \mid Q_{n,\rho_n^\alpha} \subset \Gamma] P(Q_{n,\rho_n^\alpha} \subset \Gamma) \\ &\quad + \mathbb{E}_n[G_n^{(1)} \mid Q_{n,\rho_n^\alpha} \setminus \Gamma \neq \emptyset] (1 - P(Q_{n,\rho_n^\alpha} \subset \Gamma)) \\ &\leq 0 + \mathbb{E}_n[G_n^{(1)} \mid Q_{n,\rho_n^\alpha} \setminus \Gamma \neq \emptyset] (1 - \alpha) \leq \mu(Q_{n,\rho_n^\alpha}) (1 - \alpha). \end{aligned}$$

\square