Backward Importance Sampling for Online Estimation of State Space Models

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1. Introduction

Latent data models are widely used in time series and sequential data analysis across a wide range of applied science and engineering domains such as movement ecology (Michelot et al. 2016), energy consumption modeling (Candanedo et al. 2017), genomics (Yau et al. 2011; Gassiat et al. 2016; Wang et al. 2017), target tracking (Särkkä et al. 2007), enhancement and segmentation of speech and audio signals (Rabiner 1989), see also (Särkkä 2013; Douc et al. 2014; Zucchini et al. 2017) and the numerous references therein. Performing Maximum Likelihood Estimation (MLE) for instance with the Expectation Maximization (EM) algorithm (Dempster et al. 1977) or a stochastic gradient ascent (Cappé et al. (2005) in the case of HMMs is a challenging task. Both approaches involve conditional distributions of sequences of hidden states given the observation record (the smoothing distribution), which are not available explicitly.

Markov chain Monte Carlo (MCMC) and sequential Monte Carlo (SMC) methods (also known as particle filters or smoothers) are widespread solutions to propose consistent estimators of such distributions. Among SMC methods, algorithms have been designed in the last decades to solve the smoothing problem, such as the Forward Filtering Backward Simulation algorithm (Douc et al. 2011) or two-filter based approaches (Briers et al. 2010; Fearnhead et al. 2010; Nguyen et al. 2017). These approaches, which come with strong theoretical guarantees (Del Moral et al. 2010; Douc et al. 2011; Dubarry and Le Corff 2013; Gerber and Chopin 2017), require the time horizon and all observations to be available to initialize a backward information filter, and, thus, perform the smoothing. The particle-based rapid incremental smoother (Olsson and Westerborn 2017) is an online version of forward-backward procedures, specifically designed to approximate conditional expectations of additive functionals. This algorithm relies on a backward sampling step performed on the fly thanks to an acceptance rejection procedure. This online smoother was proven to be strongly consistent, asymptotically normal, and with a control of the asymptotic variance, when it is performed together with the vanilla bootstrap filter (Gordon et al. 1993). In Olsson and Alenlöv (2020), the authors show how this algorithm can be used to perform recursive maximum likelihood estimation in the context of Partially Observed Diffusion process (POD), and in the case of a bidimensional partially observed stochastic Lotka-Volterra model. Supplementary materials for this article are available online.

This article proposes a new Sequential Monte Carlo algorithm to perform online estimation in the context of state space models when either the transition density of the latent state or the conditional likelihood of an observation given a state is intractable. In this setting, obtaining low variance estimators of expectations under the posterior distributions of the unobserved states given the observations is a challenging task. Following recent theoretical results for pseudo-marginal sequential Monte Carlo smoothers, a pseudo-marginal backward importance sampling step is introduced to estimate such expectations. This new step allows to reduce very significantly the computational time of the existing numerical solutions based on an acceptance-rejection procedure for similar performance, and to broaden the class of eligible models for such methods. For instance, in the context of multivariate stochastic differential equations, the proposed algorithm makes use of unbiased estimates of the unknown transition densities under much weaker assumptions than most standard alternatives. The performance of this estimator is assessed for high-dimensional discrete-time latent data models, for recursive maximum likelihood estimation in the context of Partially Observed Diffusion process (POD), and in the case of a bidimensional partially observed stochastic Lotka-Volterra model. Supplementary materials for this article are available online.

ABSTRACT

This article proposes a new Sequential Monte Carlo algorithm to perform online estimation in the context of state space models when either the transition density of the latent state or the conditional likelihood of an observation given a state is intractable. In this setting, obtaining low variance estimators of expectations under the posterior distributions of the unobserved states given the observations is a challenging task. Following recent theoretical results for pseudo-marginal sequential Monte Carlo smoothers, a pseudo-marginal backward importance sampling step is introduced to estimate such expectations. This new step allows to reduce very significantly the computational time of the existing numerical solutions based on an acceptance-rejection procedure for similar performance, and to broaden the class of eligible models for such methods. For instance, in the context of multivariate stochastic differential equations, the proposed algorithm makes use of unbiased estimates of the unknown transition densities under much weaker assumptions than most standard alternatives. The performance of this estimator is assessed for high-dimensional discrete-time latent data models, for recursive maximum likelihood estimation in the context of Partially Observed Diffusion process (POD), and in the case of a bidimensional partially observed stochastic Lotka-Volterra model. Supplementary materials for this article are available online.

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possible to implement importance sampling and filtering recursions, when the unavailable importance weights are replaced by random estimators. Standard data augmentation schemes were then used to extend this random-weight particle filter to provide new inference procedures for instance for partially observed diffusion models (Yonekura and Beskos 2020).

More recently, the online algorithm of Olsson and Westerborn (2017) was extended to this setting for partially observed diffusion processes by Gloaguen et al. (2018). Then, Gloaguen et al. (2022) introduced a pseudo-marginal online smoother to approximate conditional expectations of additive functionals of the hidden states in a very general setting: the user can only evaluate (possibly biased) approximations of the transition density and of the marginal conditional likelihood. The online algorithm of Gloaguen et al. (2022) may be used to approximate expectations of additive functionals under the smoothing distributions by processing the data stream online. However, as with the PaRIS algorithm, when using this pseudo-marginal approach where transition densities are intractable, the user needs to sample exactly from the associated pseudo-marginal backward kernel. This step is again done by rejection sampling, and therefore requires that the estimate of the transition density and of the marginal conditional likelihood are almost surely positive and upper bounded. In practice, these assumptions are very restrictive. For instance, in the context of diffusion processes, they narrow the possible models to the class of diffusions satisfying the Exact algorithm conditions of Beskos et al. (2006), for which General Poisson Estimators (GPEs) (Fearnhead et al. 2008) already lead to eligible unbiased estimators.

In this article, a new procedure is introduced to replace the backward acceptance–rejection step of the PaRIS and the pseudo-marginal PaRIS algorithms by a backward importance sampling estimate. It leads to a smoothing algorithm that only requires an almost surely positive estimator of the unknown transition or observation density, and therefore extends widely the class of models for which these online smoothers can be designed. In the general case where only signed estimates can be obtained, we propose to use Wald’s trick to ensure positivity. In the context of partially observed diffusion processes, for instance, we show that combining Wald’s trick to the importance sampling estimators of Andersson and Kohatsu-Higa (2017) and Fearnhead et al. (2017) leads to a very generic algorithm that can be applied to a wide class of models, for which no low variance smoother existed so far.

The article is organized as follows. Section 2 displays the latent data models and the main objectives considered in this article. Then, Section 3 details online pseudo marginal sequential Monte Carlo algorithms and Section 4 our proposed algorithm. Section 5 and Section 6 provide numerical experiments to illustrate the performance of our approach. The empirical results of this section can be summarized as follows.

- We show that the proposed approach outperforms the existing acceptance rejection method in terms of computational efficiency (Section 5.2).
- We show how the proposed method allows for efficient online recursive maximum likelihood estimation in the context of partially observed diffusion processes (Section 5.3).
- In the context of the pseudo-marginal approach, we extend the use of Wald’s trick to the backward kernel, and therefore show that our approach can be used in cases where the estimators of the unknown densities are not positive by construction.
- We perform sequential Monte Carlo smoothing in models for which no solutions were proposed in the literature to the best of our knowledge, such as multivariate partially observed diffusion processes.

## 2. Model and Objectives

Let \( \theta \) be a parameter lying in a set \( \Theta \subset \mathbb{R}^d \) and consider a state space model where the hidden Markov chain in \( \mathbb{R}^d \) is denoted by \((X_k)_{k \geq 0}\). The distribution of \( X_0 \) has density \( \chi \) with respect to the Lebesgue measure and for all \( 0 \leq k \leq n-1 \), the conditional distribution of \( X_{k+1} \) given \( X_k \) has density \( q_{k+1,\theta}(X_{k+1}, \cdot) \), where \( a_{uv} \) is a short-hand notation for \((a_u, \ldots, a_v)\). It is assumed that this state is partially observed through an observation process \((Y_k)_{0 \leq k \leq n}\) taking values in \( \mathbb{R}^m \). For all \( 0 \leq k \leq n \), the distribution of \( Y_k \) given \( X_{0:n} \) depends on \( X_k \) only and has density \( g_{k,\theta}(X_k, \cdot) \) with respect to the Lebesgue measure. In this context, for all pairs of indices \( 0 \leq k_1 \leq k_2 \leq n \), we define the joint smoothing distribution as the conditional law of \( X_{k_1:k_2} \) given \( Y_{0:n} \). In this framework, the likelihood of the observations is

\[
L_{n,\theta}(Y_{0:n}) = \int \chi(x_0)g_{0,\theta}(x_0, Y_0) \prod_{k=0}^{n-1} \ell_{k,\theta}(x_k, x_{k+1}) \, dx_{0:n},
\]

where, for all \( 0 \leq k \leq n \) and all \( \theta \in \Theta \),

\[
\ell_{k,\theta}(x_k, x_{k+1}) = q_{k+1,\theta}(x_k, x_{k+1}, Y_k)g_{k+1,\theta}(x_{k+1}, Y_{k+1}).
\]

In a large variety of situations, (1) cannot be evaluated pointwise (see for instance the model of Section 5.2), and we assume in this article that we can compute an estimate of this quantity (see assumption H1 in Section 3). Note that to avoid future cumbersome expressions, the dependency of the key quantity \( \ell_{k,\theta} \) on the observations is implicit, as we always work conditionally to the observations. For the same reason, we may write \( g_{k,\theta}(x_k) \) instead of \( g_{k,\theta}(X_k, Y_k) \) to simplify some notations.

In this article, we propose an algorithm to compute smoothing expectations of additive functionals. Namely, we aim at computing \( E_\theta [h_{0:n}(X_{0:n}) | Y_{0:n}] \), where \( E_\theta \) is the expectation under the law of the hidden Markov model parameterized by \( \theta \), and where \( h_{0:n} \) is an additive functional, that is, a function from \( \mathbb{R}^d \times (n+1) \) to \( \mathbb{R}^d \) satisfying

\[
h_{0:n} : x_{0:n} \mapsto \sum_{k=0}^{n-1} \tilde{h}_k(x_k, x_{k+1}),
\]

where \( \tilde{h}_k : \mathbb{R}^d \times \mathbb{R}^d \rightarrow \mathbb{R}^d \). Such expectations are the keystones of many common inference problems in state space models.
**Example 1 (State estimation.).** Suppose that $\theta$ is known, a common objective is to recover the underlying signal $X_k$, for some index $0 \leq k^* \leq n$ given the observations $Y_{0:n}$. A standard estimator is $E_{\theta} [X_k | Y_{0:n}]$, which is a particular instance of our problem with $\tilde{h}_k(x_k, x_{k+1}) = x_k$ if $k = k^*$ and 0 otherwise.

**Example 2 (EM algorithm.)**. In the usual case when $\theta$ is unknown, the maximum likelihood estimator is $\hat{\theta}_n = \arg\max_{\theta \in \Theta} L_n(\theta; Y_{0:n})$. Expectation Maximization based algorithms (Dempster et al. 1977) are appealing solutions to obtain an estimator of $\hat{\theta}_n$. The pivotal concept of the EM algorithm is that the intermediate quantity defined by

$$
\theta \mapsto Q(\theta, \theta') = \mathbb{E}_{\theta'} \left[ \sum_{k=0}^{n-1} \log \ell_{k,\theta}(X_k, X_{k+1}) \bigg| Y_{0:n} \right]
$$

may be used as a surrogate for $L_n(\theta; Y_{0:n})$ in the maximization procedure. Again, this inference setting is a special case of our framework where $\tilde{h}_k(x_k, x_{k+1}) = \log \ell_{k,\theta}(x_k, x_{k+1})$.

**Example 3 (Fisher’s identity and online gradient ascent.).** An alternative to the EM algorithm is to maximize the log-likelihood through gradient based methods. Indeed, in state space models, under some regularity conditions, the gradient of the log-likelihood can be obtained using Fisher’s identity:

$$
\nabla_{\theta} \log L_n(\theta; Y_{0:n}) = \mathbb{E}_{\theta} \left[ \sum_{k=0}^{n-1} \nabla_{\theta} \log \ell_{k,\theta}(x_k, x_{k+1}) \bigg| Y_{0:n} \right],
$$

which relies on the conditional expectation of an additive functional. It has been noted (see Cappe et al. 2005, chap. 10, or Olsson and Alénlöv 2020), how this identity, coupled with the smoothing recursions of Section 3.3, can lead to an online gradient ascent, that provides an online estimate of the MLE. An extension of this method is illustrated in Section 5.3 in a challenging setting where the transition density, and therefore (1), cannot be evaluated.

### 3. Online Sequential Monte Carlo Smoothing

#### 3.1. Backward Statistic for Online Smoothing

In this section, when there is no possible confusion, the parameter $\theta$ is dropped from the notations for a better clarity. For all pairs of integers $0 \leq k \leq k' \leq n$, and all measurable functions $h$ on $\mathbb{R}^d \times (k' - k + 1)$, the expectation with respect to the joint smoothing distribution is denoted by

$$
\phi_{k,k'}(h) := \mathbb{E} \left[ h(X_{k:k'}) \big| Y_{0:n} \right].
$$

The special case where $k = k'$ refers to the filtering distribution and we write $\phi_k = \phi_{k,k}$. A pivotal quantity to estimate $\phi_{0:n}(h_{0:n})$ is the backward statistic:

$$
T_k[h_{0:k}](X_k) = \mathbb{E} \left[ h_{0:k}(X_{0:k})X_k \big| Y_{0:k} \right], \quad 1 \leq k \leq n, \quad T_0 = 0.
$$

Note that for each $k$, $T_k[h_{0:k}](X_k)$ is a function of $X_k$, and is defined relatively to the functional of interest $h_{0:k}$. For additive functionals, this statistic satisfies the following two key identities, for all $1 \leq k \leq n$:

$$
\phi_{0:n}[h_{0:n}] = \mathbb{E} \left[ h_{0:n}(X_{0:n}) \big| Y_{0:n} \right] = \mathbb{E} \left[ \left. \mathbb{E} \left[ h_{0:n}(X_{0:n}) \big| Y_{0:n}, X_n \right] \right| Y_{0:n} \right] = \phi_n \left[ T_n[h_{0:n}] \right],
$$

$$
T_k \left[ h_{0:k} \right](X_k) = \mathbb{E} \left[ T_{k-1} \left[ h_{0:(k-1)} \right] (X_{k-1}) + \tilde{h}_{k-1} (X_{k-1}, X_k) \big| X_k, Y_{0:k-1} \right],
$$

where $\tilde{h}_{k-1}$ is the function defined in (2). Property (4), which results from the tower property for conditional expectations means that the target quantity is the filtering expectation of a well chosen statistic, while property (4) provides a recursion to compute these statistics. These two properties suggest an online procedure to solve the online smoothing problem. Starting at time 0, at each step $k$, this procedure aims at (i) computing the filtering distribution and (ii) computing the backward statistics.

Following Fearnhead et al. (2008), Olsson and Ströjby (2011), and Gloaguen et al. (2018, 2022), we do not assume that (1) can be evaluated pointwise. We assume that there exists an estimator, depending on some random variable in a general state space $(U, B(U))$ such that the following assumption holds:

**H1** For all $\theta \in \Theta$ and $k \geq 0$, there exists a Markov kernel on $(\mathbb{R}^d \times \mathbb{R}^d, B(U))$ with density $R_{k,\theta}$ with respect to a reference measure $\mu$ on a general state space $(U, B(U))$, and a positive mapping $\ell_{k,\theta}(\cdot)$ on $\mathbb{R}^d \times \mathbb{R}^d \times U$ such that, for all $(x, x') \in \mathbb{R}^d \times \mathbb{R}^d$,

$$
\int R_{k,\theta}(x, z; z) \ell_{k,\theta}(z)(x', \mu)(dz) = \ell_{k,\theta}(x, x').
$$

This framework, known as pseudo marginalisation, is based on the plug-in principle, as a pointwise estimate of $\ell_{k,\theta}(x, x')$ can be obtained by generating $\zeta$ from $R_{k,\theta}(x, x'; dz)$ and computing the statistic $\ell_{k,\theta}(\zeta)(x, x')$. Its use in Monte Carlo methods, and the related theoretical guarantees, have been studied in the context of MCMC (Andrieu and Roberts 2009), and more recently, of SMC (Gloaguen et al. 2022).

#### 3.1.1. Recursive Maximum Likelihood Estimation

An appealing application for online smoothing is the context of recursive maximum likelihood estimation, that is, where new observations are used only once to update the estimator of the unknown parameter $\theta$. Following Le Gland and Mevel (1997), the idea is to build a sequence $\{\theta_k\}_{k \geq 0}$ as follows. First, set the initial value $\theta_0$ of the parameter estimate and then, for each new observation $Y_k$, $k \geq 1$, define

$$
\theta_k = \theta_{k-1} + \gamma_k \nabla_{\theta} \ell_{\theta}(Y_k | Y_{0:k-1}),
$$

where $\ell_{\theta}(Y_k | Y_{0:k-1})$ is the log-likelihood for the new observation given all the past, and $\{\gamma_k\}_{k \geq 1}$ are positive step sizes such that $\sum_{k \geq 1} \gamma_k = \infty$ and $\sum_{k \geq 1} \gamma_k^2 < \infty$. The practical implementation of such an update relies on the following identity:

$$
\nabla_{\theta} \ell_{\theta}(Y_k | Y_{0:k-1}) = \frac{\pi_{k,\theta} \nabla_{\theta} \ell_{\theta} + \eta_{k,\theta} \eta_{k,\theta}}{\pi_{k,\theta}}
$$

where $\pi_{k,\theta} = \phi_{k,\theta} | Y_{0:k-1}$ is the predictive distribution at time $k$, that is, the law of $X_k$ given $Y_{0:k-1}$, and

$$
\eta_{k,\theta} = \phi_{0:k,\theta} | Y_{0:k-1} - \pi_{k,\theta} \phi_{0:k,\theta} | Y_{0:k-1}.
$$
with
\[ h_{0:k}(x_{0:k}) = \sum_{j=0}^{k-1} \nabla_{\theta} \log \ell_{j,\theta}(x_j, x_{j+1}). \] (8)

The signed measure \( \eta_{k,\theta} \) is known as the "tangent filter," see (Cappé et al. 2005, chap. 10), Del Moral et al. (2015) or Olsson and Alenlöv (2020). Using the tower property and the backward decomposition (5) yields
\[
\begin{align*}
\phi_{0:k|\theta|k-1}[h_{0:k|\theta|k}g_{k,\theta}] &= \mathbb{E}_{\theta_0} \left[ h_{0:k|\theta}(X_{0:k})g_{k,\theta}(X_k) | Y_{0:k-1} \right], \\
&= \mathbb{E}_{\theta} \left[ h_{0:k|\theta}(X_{0:k})X_k, Y_{0:k-1} g_{k,\theta}(X_k) | Y_{0:k-1} \right], \\
&= \mathbb{E}_{\theta} \left[ T_k h_{0:k}(X_k)g_{k,\theta}(X_k) | Y_{0:k-1} \right] = \pi_{k,\theta} \left[ T_k h_{0:k|\theta|k}g_{k,\theta} \right], \\
& \text{so that} \\
\eta_{k,\theta}[g_{k,\theta}] &= \pi_{k,\theta} \left[ (T_k h_{0:k} - \pi_{k,\theta}[T_k h_{0:k|\theta|k}]) g_{k,\theta} \right]. \quad \text{(9)}
\end{align*}
\]

It is worth noting that, in the context of this article where \( \ell_k \) cannot be evaluated pointwise, the functional (8) is unknown. In Section 5.3, we illustrate that we can plug-in an estimate of this functional instead. The rationale motivating the algorithm given by (6) relies on the following expression of the normalized log-likelihood:
\[
\frac{1}{n} \nabla_{\theta} \ell_0(Y_{0:n-1}) = \frac{1}{n} \sum_{k=0}^{n-1} \nabla_{\theta} \ell_0 \left( Y_k \mid Y_{0:k-1} \right),
\]
with the convention \( \ell_0(Y_0 | Y_{0:k-1}) = \ell_0(Y_0) \). Moreover, under strong mixing assumptions, for all \( \theta \in \Theta \), the extended process \( \{(X_n, Y_n, \pi_n, \eta_n) \}_{n \geq 0} \) is an ergodic Markov chain and for all \( \theta \in \Theta \), the normalized score \( \nabla_{\theta} \ell_0(Y_{0:n-1})/n \) converges almost surely to a limiting quantity \( \lambda(\theta, \theta_0) \) such that, under identifiability constraints, \( \lambda(\theta_0, \theta_0) = 0 \). A gradient ascent algorithm cannot be designed as the limiting function \( \theta \mapsto \lambda(\theta, \theta_0) \) is not available explicitly. However, solving the equation \( \lambda(\theta_0, \theta_0) = 0 \) may be cast into the framework of stochastic approximation to produce parameter estimates using the Robbins-Monro algorithm
\[
\theta_k = \theta_{k-1} + \gamma_k \zeta_k, \quad n \geq 0, \quad \text{(10)}
\]
where \( \zeta_k \) is a noisy observation of \( \lambda(\theta_{k-1}, \theta_0) \), equal to (7). In the case of a finite state space \( X \) the algorithm was studied in Le Gland and Mevel (1997), which also provides assumptions under which the sequence \( \{\theta_n\}_{n \geq 0} \) converges toward the parameter \( \theta_0 \) (see also Tadić (2010) for refinements).

### 3.2. Approximation of the Filtering Distribution

Let \( \{\xi^\ell_k\}_{\ell=1}^N \) be independent and identically distributed according to an instrumental proposal density \( \rho_0 \) on \( \mathbb{R}^d \) and define the importance weights \( \omega_k = \chi(\xi^\ell_k, \theta_0)g_0(\xi^\ell_k, Y_0)/\rho(\xi^\ell_k) \), \( \chi \) is the density of the distribution of \( X_0 \) as defined in Section 2. For any bounded and measurable function \( h \) defined on \( \mathbb{R}^d \), the importance sampling estimator defined as
\[
\phi^N_\ell[h] = \Omega_0^{-1} \sum_{\ell=1}^N \omega_0 h(\xi^\ell_k), \quad \text{where} \quad \Omega_0 := \sum_{\ell=1}^N \omega_0.
\]
is a consistent estimator of \( \phi_\ell[f] \). Then, for all \( \ell \geq 1 \), once the observation \( Y_k \) is available, particle filtering transforms the weighted particle sample \( \{\omega^\ell_k, \xi^\ell_k\}_{k=1}^N \) into a new weighted particle sample approximating \( \phi_\ell \). This update step is carried through in two steps, selection and mutation, using sequential importance sampling and resampling steps. New indices and particles \( \{l_k^\ell, \xi^\ell_k, \xi^\ell_k\}_{\ell=1}^N \) are simulated independently from the instrumental distribution with density on \( \{1, \ldots, N\} \times \mathbb{R}^d \times U \):\[
\omega_k(l, x, z) \propto \omega^\ell_k p_{k-1}(\xi^\ell_k, x) R_k(\xi^\ell_k, x, | z),
\]
where \( p_{k-1} \) is a Markovian transition density. In practice, this step is performed as follows:

1. Sample \( l_k^\ell \) in \( \{1, \ldots, N\} \) with probabilities proportional to \( \{\omega^\ell_k\}_{l=1}^N \) with probabilities proportional to \( \{\omega^\ell_k\}_{l=1}^N \).

2. Sample \( \xi^\ell_k \) with distribution \( p_{k-1}(\xi^\ell_k, \cdot) \) and sample \( \xi^\ell_k \) with distribution \( R_k(\xi^\ell_k, \xi^\ell_{k-1}, \cdot) \).

For any \( \ell \in \{1, \ldots, N\} \), \( \xi^\ell_k \) is associated with the importance weight defined by
\[
\omega_k^\ell = \frac{\ell_{k-1}(\xi^\ell_k, \xi^\ell_{k-1})}{p_{k-1}(\xi^\ell_k, \xi^\ell_{k-1})} \quad \text{to produce the following approximation of} \phi_k[f]:
\]
\[
\phi_k^N[f] := \Omega_k^{-1} \sum_{\ell=1}^N \omega_k^\ell f(\xi^\ell_k), \quad \text{where} \quad \Omega_k := \sum_{\ell=1}^N \omega_k^\ell.
\]
The choice of the proposal distribution \( p_{k-1} \) is a pivotal tuning step to obtain efficient estimations of the filtering distributions. This point is discussed in each example considered in Section 5.

### 3.3. Approximation of the Backward Statistics

Approximation of the backward statistics, as defined in (3), are computed recursively, for each simulated particle. The computations starts with initializing a set \( \tau_0^i = 0, \ldots, \tau_0^N = 0 \), corresponding to the values of \( T_0(\xi^\ell_0), \ldots, T_0(\xi^\ell_N) \). Then, using (5), for each \( \ell \geq 0 \), \( 1 \leq i \leq N \), the approximated statistics are updated with
\[
\tau^{i}_{k+1} = \frac{1}{N} \sum_{j=1}^N \left( l_{k+1}^{i, j} + \tilde{h}_k(\xi_k^{i, j}, \xi_{k+1}^{i, j}) \right), \quad \text{(12)}
\]
where \( \tilde{N} \geq 1 \) is a sample size which is typically small compared to \( N \) (Olsson and Westerborn 2017 typically suggests to take \( \tilde{N} = 2 \)) and where \( U_{k+1}^{i, j} \) is iid in \( \{1, \ldots, N\} \times U \) with distribution
\[
\pi_k^i(\ell, z) \propto \omega_k^\ell \xi_k^\ell(z)(\xi^\ell_k, \xi^\ell_{k+1}) R_k(\xi^\ell_k, \xi^\ell_{k+1}; z).
\]
As explained in Gloaquet et al. (2022), this recursive update requires to produce samples \( l_{k+1}^{i, j} \) distributed according to the marginal distribution of \( \pi_k^i \), referred to as the backward kernel. In practice, this requires computationally intensive sampling procedures and the only proposed practical solution can be
used in very restrictive situations. In Gloaguen et al. (2018), the authors assumed that almost surely, for all \( x, x', \ell_k(\zeta)(x, x') \geq 0 \), and that, for all \( 0 \leq k \leq n \) and \( 0 \leq i \leq N \), there exists an upper bound \( \epsilon_k^i \) such that

\[
\sup_{\zeta} \ell_k(\zeta)(\xi^i_k, \xi^i_{k+1}) \leq \epsilon_k^i .
\]

Then, if the positiveness assumption of \( \ell_k(\zeta)(x, x') \) is satisfied, sampling from the distribution \( T_{\zeta} \) is possible under assumption (13), as, for all \((i, z) \in \{1, \ldots, N\} \times U_k\),

\[
\omega^i_k \ell_k(z)(\xi^i_k, \xi^i_{k+1})R_k(\xi^i_k, \xi^i_{k+1}; z) \leq \epsilon_k^i \omega^i_k R_k(\xi^i_k, \xi^i_{k+1}; z) .
\]

Therefore, the following accept–reject method may be used to sample from \( \eta^i_k \).

1. A candidate \((f^*, \zeta^*)\) is sampled in \( \{1, \ldots, N\} \times U \) as follows:
   
   (a) \( f^* \) is sampled with probabilities proportional to \( (\omega^i_k)_{i=1}^N \);
   
   (b) \( \zeta^* \) is sampled independently with distribution \( R_k(\xi^i_k, \xi^i_{k+1}; \zeta^*) \).

2. \((f^*, \zeta^*)\) is accepted with probability \( \ell_k(\zeta^*)(\xi^i_k, \xi^i_{k+1})/\epsilon^i_k \).

Upon acceptance, \( f_{k+1}^i = f^* \).

This algorithm is the only online SMC smoother proposed in the literature with theoretical guarantees when no closed-form expressions of the transition densities and the conditional likelihood of the observations are available, assuming that the user can only evaluate approximations of these densities. This pseudo-marginal particle smoothing algorithm requires that the backward sampling step generates samples exactly according to \( \eta^i_k \). However, it relies on the key assumptions of the positiveness of \( \ell_k(\zeta)(x, x') \) and (13) which are rather restrictive (especially the second one), and would not be satisfied in practice for a lot of problems (see for instance in Section 5). In Section 4, we propose an alternative to this step to obtain a computationally efficient pseudo-marginal smoother in a much wider range of applications for which such assumptions do not hold.

### 3.3.1. Approximations for Recursive MLE

In the case of recursive MLE, it is required to approximate the key quantity (7). A particle filter, can be used to compute the following sequential Monte Carlo approximations:

\[
\pi_k^N[g_k, \theta] = \frac{1}{N} \sum_{i=1}^N \nabla \ell^i_k g_k(\xi^i_k) , \quad \pi_k^N[V_\theta g_k, \theta] = \frac{1}{N} \sum_{i=1}^N \nabla \ell^i_k g_k(\xi^i_k) .
\]

In addition, the tangent filter can be approximated using a backward sampling procedure, based on the backward statistic associated with the functional (8):

\[
\eta_k^N[g_k, \theta] = \frac{1}{N} \sum_{i=1}^N \tau^i_k g_k(\xi^i_k) - \left( \frac{1}{N} \sum_{i=1}^N \tau^i_k \right) \left( \frac{1}{N} \sum_{i=1}^N g_k(\xi^i_k) \right) .
\]

Plugging these estimates in (7) allows to perform the online recursive algorithm.

### 4. Pseudo-Marginal Backward Importance Sampling

#### 4.1. Positive Estimates

In this section, we propose to use Wald’s identity for martingales to obtain an estimator which is guaranteed to be positive. This step is not required if \( \ell_k(z)(x, x') \) is positive by construction but this is not necessarily true. Wald’s trick was for instance applied in Fearnhead et al. (2010) to solve the filtering problem in the context of Poisson based estimators for partially observed diffusions. Our estimator is defined up to an unknown constant of proportionality, which is removed when the importance weights are normalized in (15). This approach, rather than setting negative weights to 0, which would lead to a biased estimate, uses extra simulation to obtain positiveness. This is done while ensuring that the weights remain unbiased up to a common constant of proportionality.

**Particle filtering weights.** For all \( k \geq 0 \), the Wald-based random weight particle filtering proceeds as follows.

1. For all \( 1 \leq i \leq N \), sample a new particle as described in Section 3.2.

   (a) Sample \( I^i_k \) in \( \{1, \ldots, N\} \) with probabilities proportional to \( (\omega^i_{k-1})_{1 \leq j \leq N} \).

   (b) Sample \( \xi^i_k \) with distribution \( p_{k-1}(\xi^i_{k-1}, \cdot) \).

2. For all \( 1 \leq i \leq N \), set \( \omega^i_k = 0 \).

3. While there exists \( i_* \in \{1, \ldots, N\} \) such that \( \omega^i_{k-1} \leq 0 \), for all \( 1 \leq i \leq N \), sample \( \xi^i_k \) with distribution \( R_k(\xi^i_{k-1}, \xi^i_{k-1}, \cdot) \), that is, compute an estimator of the transition density, and set

\[
\omega^i_k = \omega^i_k + \frac{\ell_k(\xi^i_k)(\xi^i_{k-1}, \xi^i_{k-1})}{p_{k-1}(\xi^i_{k-1}, \xi^i_{k-1})} .
\]

We aim now at updating the backward statistics \( \tau^i_{k+1}, 1 \leq i \leq N \) using an importance sampling step. Therefore, we introduce the following extension of Wald’s trick importance sampling to the online smoothing setting of this article.

**Backward simulation weights.** For all \( 1 \leq i \leq N \), the backward importance sampling step proceeds then as follows.

1. For all \( 1 \leq j \leq N \), sample \( J^{(ij)}_k \) in \( \{1, \ldots, N\} \) with probabilities proportional to \( (\omega^j_{k-1})_{i=1}^N \).

2. For all \( 1 \leq j \leq N \), set \( \sigma^{(ij)}_k = 0 \).

3. While there exist \( j_* \in \{1, \ldots, N\} \) such that \( \sigma^{(ij)}_k \leq 0 \), for all \( 1 \leq j \leq N \), sample \( \zeta^{(ij)}_k \) with distribution \( R_k(\xi^{(ij)}_{k+1}, \xi^{(ij)}_{k+1}, \cdot) \) and set

\[
\sigma^{(ij)}_k = \sigma^{(ij)}_k + \ell_k(\zeta^{(ij)}_k)(\xi^{(ij)}_{k+1}, \xi^{(ij)}_{k+1}) .
\]
4.2. AR-Free Online Smoothing

Without any additional assumption, the statistics are then updated recursively as follows: for all \( 1 \leq i \leq N \),

\[
\tau^{(i)}_{k+1} = \sum_{j=1}^{\tilde{N}} \mathcal{W}^{(i)}_{k} \left( \tau^{(i)}_{k+1} + \tilde{h}_{k} \left( \tilde{X}^{(i)}_{k+1}, \tilde{X}^{(i)}_{k+1} \right) \right),
\]

(15)

where \( \mathcal{W}^{(i)}_{k} \), \( 1 \leq j \leq \tilde{N} \) are computed using the pseudo-marginal smoothing technique combined with Wald’s identity and \( \mathcal{W}^{(i)}_{k} = \sum_{j=1}^{\tilde{N}} \mathcal{W}^{(i)}_{k} \). Then, the estimator of the conditional expectation of the additive functional \( \phi_{0:n|n}[h_{0:n}] \) is set as

\[
\phi^{N,IS}_{0:n|n}[h_{0:n}] := \sum_{i=1}^{N} \frac{\tilde{o}^{(i)}_{n}}{\Omega_{n}} \tilde{r}^{(i)}_{n}.
\]

5. Application to Smoothing Expectations and Score Estimation

5.1. Recurrent Neural Networks

Recurrent Neural Networks (RNNs) were first introduced in Mozer (1989) to model time series using a hidden context state. Such deep learning models are appealing to describe short time dependencies, and RNN extensions ( Hochreiter and Schmidhuber 1997; Cho et al. 2014 ) are since then widely used in practice, see for instance Mikolov et al. (2010) and Sutskever et al. (2011, 2014). In this section, we propose a general state space model based on a vanilla RNN architecture, as follows. The hidden state is initialized as \( X_{0} \sim \mathcal{N}(0, \Sigma) \) and for all \( k \geq 1 \),

\[
X_{k} = \tanh(W_{1} Y_{k-1} + W_{2} X_{k-1} + b + \eta_{k}) \quad \text{and} \quad Y_{k} = W_{3} X_{k} + c + \varepsilon_{k},
\]

where \( W_{1}, W_{2}, W_{3}, \) and \( b \) and \( c \) are the weight matrices and bias, \( \Sigma \) is an unknown covariance matrix and \((\eta_{k})_{k \geq 1}\) and \((\varepsilon_{k})_{k \geq 1}\) are independent Gaussian random variables with covariance matrices \( Q \) and \( R \). In this experiment, we show that the proposed algorithm can be used in this setting which does not fit the usual assumptions of hidden Markov models. While we here focus on a simple vanilla one-layer recurrent network, such a general state space model could be extended to multi-layer RNN architectures by considering noisy state dynamics in each hidden layer, and to RNN variants such as Long Short Term Memory ( Hochreiter and Schmidhuber 1997) and Gated Recurrent Unit ( GRU ) ( Cho et al. 2014).

To generate a synthetic sequence of states and observations \( (X_{0:n}, Y_{0:n}) \) from this stochastic RNN, we considered diagonal covariance matrices \( \Sigma, Q, \) and \( R \), with the same variance along all dimensions equal to 0.1. To obtain weights and biases values corresponding to realistic data, we trained a deterministic one-layer RNN on 20,000 samples of a weather time-series dataset available online. In such a setting, the observations \( Y_{0:n} \) consist in a sequence of 4D vectors (originally temperature, air pressure, air humidity, air density). Following the classical RNN framework, the sequence of hidden states \( X_{0:n} \) is usually made of higher-dimensional vectors; the experiments were performed for two RNNs of respective hidden dimension 32 and

64. After this training part, we sampled \( (X_{0:n}, Y_{0:n}) \) according to the model. For each RNN, a single sequence of hidden states and observations was simulated with a total length of 200 time steps.

From this general state space model based on a stochastic RNN, in the context of the state estimation problem, we are interested in using particle smoothing algorithms to estimate two smoothing expectations, respectively \( \mathbb{E}[X_{0}|Y_{0:n}] \), and \( \mathbb{E}[\sum_{k=0}^{n} X_{k}|Y_{0:n}] \). In our context of online estimation, the evaluation was made for \( n = 49 \) (sequence truncated at 50 observations), \( n = 99 \) (sequence truncated at 100 observations), and \( n = 199 \) (full sequence). The Monte Carlo estimate of these quantities is referred to with the hat symbol: \( \hat{\mathbb{E}}[\cdot] \).

In the following, the performance of our algorithm was compared to the classical Poor Man’s smoother. The Poor Man’s smoother (also known as the path-space smoother) estimates the joint smoothing distribution using the ancestral lines of each particle at time \( n \); see for instance Douc et al. (2014) for discussions on the path degeneracy issue. For the backward IS smoother, we use \( N = 1000 \) particles for the bootstrap filter and \( \tilde{N} = 32 \) backward samples (see Section 5.2 and Figure 3 for the choice of \( \tilde{N} \) the number of particles \( N \)). For the Poor Man’s smoother \( N = 3000 \) particles were used, which yields a similar computational cost than the backward IS smoother. One interesting aspect of applying the backward IS smoother (BIS) on neural network architectures is the parallelization abilities of such algorithm: the loop over the backward samples in the backward step of the BIS is easily parallelizable, while the Poor Man’s smoother requires to store the full past trajectories of the particles.

Table 1 displays the results from performing 100 runs for each smoothing algorithm. The performance metric is the classical Mean Squared Error (MSE), which is approximated with the empirical mean over the 100 runs. The backward IS smoother outperforms the Poor Man’s smoother when estimating both quantities. This is also illustrated in Figure 1, displaying for the stochastic RNN of dimension 64 the MSE over 100 runs of \( \mathbb{E}[X_{k}|Y_{0:199}] \), for \( k \in \{0, \ldots, 199\} \): the backward IS smoother has a significantly smaller MSE than the Poor Man’s smoother for all observations that are recorded far in the past (in our example, for all \( k \leq 150 \)). Moreover, the table also shows that while the backward IS’s MSE tends to stay stable for all given \( n \) (49, 99, and 199), as expected the Poor Man’s estimation is less accurate for a longer sequence of observations, with the MSE increasing as \( n \) increases.

Table 1. Empirical estimation of \( \mathbb{E}[X_{0} - \hat{X}(X_{0}|Y_{0:n})]^{2} \) and \( \mathbb{E}[\sum_{k=0}^{n} X_{k} - \hat{X}(X_{k}|Y_{0:n})]^{2}/(n+1) \) for the Poor man’s smoother (PMS) and the Backward IS smoother, for states \( X_{0:n} \) and observations \( Y_{0:n} \) generated with stochastic RNNs of dimension 32 and 64.

| RNN dim | 32 | 64 |
|---------|----|----|
| PMS     |    |    |
| Backward IS |    |    |
| MSE     | 0.2147 0.1997 0.1969 0.1649 | 0.2135 0.1997 0.1914 0.1649 |
| MSE     | 0.2551 0.2056 0.1822 0.1427 | 0.2531 0.2189 0.1775 0.1519 |
| MSE     | 0.2135 0.1997 0.1914 0.1649 | 0.2018 0.1997 0.1707 0.1650 |
| MSE     | 0.2307 0.2147 0.1633 0.1589 | 0.2147 0.1997 0.1707 0.1650 |

NOTE: We consider an online estimation of a sequence of 200 observations, truncated at timestep \( n = 49 \) (50 timesteps), \( n = 99 \) (100 timesteps) and the full sequence (\( n = 199 \)).

1 https://www.bgc-jena.mpg.de/wetter
5.2. One Dimensional Diffusion Process: The Sine Model

This section investigates the performance of the proposed algorithm to compute expectations under the smoothing distributions in a context where alternatives are available for comparison. Consider the Sine model where \((X_t)_{t \geq 0}\) is assumed to be a weak solution to

\[
\mathrm{d}X_t = \sin(X_t - \theta) \, \mathrm{d}t + \mathrm{d}W_t, \quad X_0 = x_0.
\]

This simple model has no explicit transition density, however, a General Poisson estimator which satisfies (13) can be computed by simulating Brownian bridges, (see Beskos et al. 2006). Therefore, the backward importance sampling technique proposed in this article can be compared to the usual acceptance–rejection algorithm described in Section 3.3. For this simple comparison, observations are received at evenly spaced times \(t_0 = 0, \ldots, t_{10} = 5\) from the model

\[
Y_k = X_{t_k} + \varepsilon_k, \quad 0 \leq k \leq n = 10,
\]

where \((\varepsilon_k)_{0 \leq k \leq 10}\) are iid Gaussian random variables with mean 0 and variance 1. In this experiment \(\theta = \pi/4\). The proposal distribution \(p_k\) for the particle filtering approximation is chosen as the following approximation of the optimal filter:

\[
p_k(x_k, x_{k+1}) \propto q^\text{Eul}_{k+1}(x_k, x_{k+1}) g_{k+1}(x_{k+1}, Y_{k+1}),
\]

where \(q^\text{Eul}_{k+1}\) is the probability density function of the law of \(Y_k\) given \(X_k\) that is, of a Gaussian random variable with mean \(X_k\) and variance 1. As the observation model is linear and Gaussian, the proposal distribution is therefore Gaussian with explicit mean and variance.

In this first experiment, particles are used to solve the state estimation problem for the first observation that is, to compute an estimate of \(E[X_0 | Y_{0:n}]\). Figure 2 displays the computational complexity and the estimation of the posterior mean with the accept-reject algorithm and the proposed backward sampling technique as a function of \(\tilde{N}\). In this setting, \(N = 100\), and each unbiased estimate is computed using 30 Monte Carlo replicates.

For \(\tilde{N} = 2\) (which is the recommended value for the PaRIS algorithm, see Olsson and Westerborn 2017), our estimate shows a bias, which is not surprising, as it is based on a biased normalized importance sampling step. However, this bias quickly vanishes for \(\tilde{N} \geq 10\). Interestingly, our method comes with a drastic (of a factor of 10) reduction of computational time. The vanishing of the bias might induce more backward sampling, but this remains much faster than the acceptance rejection method with \(\tilde{N} = 2\).

Then, the same estimation was performed (on the same dataset) for \(N\) varying from 50 to 2000. In this context, \(\tilde{N}\) was set to 2 for the AR method. To have an empirical intuition of how \(\tilde{N}\) must vary with \(N\) for our algorithm, the backward importance sampling is applied with \(\tilde{N} = N^{0.5}, N^{0.6}\) and \(N/10\) (as this last value was sufficient in the first experiment to avoid any bias).

The results are shown in Figure 3. A small bias might appear for \(N = 2000\) and \(\tilde{N} = 45\) (≈ 2000^{0.5}), but no bias is visible.
for \(N^{0.6}\) and \(N/10\). As expected, the gain in time, compared to the state of the art algorithm, remains important (even if it decreases as \(N\) increases). It is worth noting that the variance of the computational time is greatly reduced compared to the AR technique.

### 5.3. Recursive Maximum Likelihood Estimation in the Sine Model

Online recursive maximum likelihood using pseudo marginal SMC is illustrated for the same Sine model. As mentioned, a GPE estimator of the transition density can be computed. Following the idea of this computation, it is possible to obtain an unbiased estimate of the gradient of the log-transition density and thus compute and unbiased estimate of the key quantity given in (8).

To the best of our knowledge, this estimator is new, and given in Appendix A. Using the Exact algorithm of Beskos et al. (2006) a dataset of 5000 points (displayed in Figure 4), was simulated using the true parameter \(\theta_0 = \pi/4\). As in the previous section, particle smoothing was performed, using the same particle filter, with \(N = 100\) particles and \(\tilde{N} = 10\) backward samples in our backward importance sampling procedure. In this setup, we explore three key features of our estimator.

**Sensitivity to the starting point \(\hat{\theta}_0\).** The inference procedure was performed on the same dataset from 50 different starting points uniformly chosen in \((0, 2\pi)\). The gradient step size \(\gamma_k\) of (10) was chosen to be constant (and equal to \(0.5\)) for the first 300 time steps, and then decreasing with a rate proportional to \(k^{-0.6}\). Results are given in Figure 4. There is no sensitivity to the starting point of the algorithm, and after a couple of hundred observations, the estimates all concentrate around the true value. As the gradient step size decreases, the estimates stay around the true value following autocorrelated patterns that are common to all trajectories.

**Asymptotic normality.** The inference procedure was performed on 50 different datasets simulated with the same \(\theta_0\). The 50 estimates were obtained from the same starting point (fixed to \(\theta_0\), as Figure 4 shows no sensitivity to the starting point). Figure 5 shows the results for the raw and the averaged estimates. The averaged estimates \((\tilde{\theta}_k)_{k \geq 0}\) consist in averaging the values produced by the estimation procedure after a burning phase of \(n_0\) time steps (here \(n_0 = 300\) time steps). This procedure allows to obtain an estimator whose convergence rate does not depend on the step sizes chosen by the user, see Polyak and Juditsky (1992) and Kushner and Yin (1997). For all \(0 \leq k \leq n_0\), \(\tilde{\theta}_k = \hat{\theta}_k\) and for all \(k > n_0\),

\[
\tilde{\theta}_k = \frac{1}{k - n_0} \sum_{j=n_0+1}^{k} \hat{\theta}_j.
\]
The estimated distribution of the final estimates seems to be Gaussian, centered around the true value. This conjecture, which would extend the asymptotic normality obtained in Gloaguen et al. (2022) for the original pseudo-marginal PaRIS, should be proven in future works.

Step size influence. To illustrate the influence of the gradient stepsizes, different settings are considered. In each scenario, the sequence \( (\gamma_k)_{k \geq 0} \) is given by

\[
\gamma_k = \gamma_0 \mathbb{1}_{k \leq n_0} + \gamma_0 \left( \frac{k - n_0}{k - n_0 + 1} \right) \mathbb{1}_{k > n_0},
\]

where \( \gamma_0 = 0.5 \). In this experiment \( \kappa \in \{0.5, 0.6, 0.7, 0.8, 0.9, 1\} \). The results are shown in Figure 6. As expected, the raw estimator shows different rates of convergence depending on \( \kappa \), whereas the averaged estimator has the same behavior in all cases.

6. Applications to Real Data

We now focus on real data applications, and specifically on abundance estimation of predator-prey populations in a context of noisy observations. The Lotka-Volterra model is specified in continuous time, as fully discussed in Hening and Nguyen (2018). This challenging context aims at illustrating how our method performs in a context where state-of-the-art unbiased methods, such as the Exact algorithms of Beskos et al. (2006), cannot be performed. We also propose in the supplementary materials an additional real data application using a stochastic volatility model.

In a continuous time Lotka-Volterra framework, we assume that \( (X_t)_{t \geq 0} \) is a bivariate process giving the actual abundance of predators and prey is assumed to follow the following SDE:

\[
dX_t = \alpha_\theta(X_t)dt + \begin{pmatrix} X_1(t) \\ 0 \\ X_2(t) \end{pmatrix} \Gamma dW_t,
\]

where \( W_t \) is a vector of independent standard Wiener processes, \( \Gamma \) a \( 2 \times 2 \) matrix, and for \( \kappa = (x_1, x_2)^T \):

\[
\alpha_\theta(\kappa) = \begin{pmatrix} x_1(a_{10} - a_{11}x_1 - a_{12}x_2) \\ x_2(-a_{20} + a_{21}x_1 - a_{22}x_2) \end{pmatrix}.
\]

In this context, the unknown parameter to be estimated is \( \theta = (a_{10}, a_{11}, a_{12}, a_{20}, a_{21}, a_{22}, \Gamma) \). The observation model follows a widespread framework in ecology where the abundance of preys and predators are observed through some abundance index at discrete times \( t_0, \ldots, t_n \):

\[
Y_{t_k} = \begin{pmatrix} c_1X_1(t_k)e^{\epsilon_{1k}} \\ c_2X_2(t_k)e^{\epsilon_{2k}} \end{pmatrix},
\]

where \( c = (c_1, c_2)^T \) is known (the observed fraction of the population) and \( \{\epsilon_{1k}, \epsilon_{2k}\}_{1 \leq k \leq n} \) are iid random variables distributed as \( N_2(-\text{diag} \Sigma/2, \Sigma) \) where \( \Sigma \) is an unknown
A. General Poisson Estimators

Let \( (X_t)_{t \geq 0} \) be defined as a weak solution to the following Stochastic Differential Equation (SDE) in \( \mathbb{R}^d \):

\[
X_0 = x_0 \quad \text{and} \quad dX_t = \alpha_\theta(X_t)dt + dW_t, \tag{20}
\]

where \((W_t)_{t \geq 0}\) is a standard Brownian motion, \(\alpha_\theta : \mathbb{R}^d \to \mathbb{R}^d\) is the drift function. The inference procedure presented in this article is applied in the case where the solution to (20) is supposed to be partially observed at times \(t_0 = 0, \ldots, t_n\), for a given \(n \geq 1\), through an observation process \((Y_k)_{0 \leq k \leq n}\) taking values in \(\mathbb{R}^m\). For all \(0 \leq k \leq n\), the distribution of \(Y_k\) given \((X_t)_{t \geq 0}\) depends on \(X_k = X_{t_k}\) only and has density \(g_{k,\theta}\) with respect to \(\nu\). The distribution of \(X_0\) has density \(\chi\) with respect to \(\mu\) and for all \(0 \leq k \leq n-1\), the conditional distribution of \(X_{k+1}\) given \((X_t)_{0 \leq t \leq k}\) has density \(g_{k+1,\theta}(X_{k+1})\) with respect to \(\mu\).

A.1. Unbiased Estimators of the Transition Densities

The algorithm described above strongly relies on assumption H1. In the context of SDEs, when \(g_{k+1,\theta}\) is available explicitly, this boils down to finding an unbiased estimate \(\hat{q}_{k+1,\theta}(x, y)\) of \(g_{k+1,\theta}(x, y)\) and defining

\[
\ell_{k,\theta}(x, y) = \hat{q}_{k+1,\theta}(x, y)g_{k+1,\theta}(x_{k+1}, Y_{k+1}).
\]

A.2. General Poisson Estimators

In Olsson and Ströjby (2011) and Gloaeken et al. (2018), General Poisson Estimators (GPEs) are used to obtain an unbiased estimate of the transition density. However, designing such estimators requires three strong assumptions (Beskos et al. 2006).
1. The diffusion defined by (20) can be transformed into a unit diffusion through the Lamperti transform, with drift function \( \tilde{a}_\theta(x) \).

2. The drift of this unit diffusion can be expressed as the gradient of a potential function, that is, there exists a twice differentiable function \( A_\theta : \mathbb{R}^d \to \mathbb{R} \) such that \( \tilde{a}_\theta = \nabla_x A_\theta \).

3. The function \( x \mapsto (\|a_\theta(x)\|^2_+ + \Delta A_\theta(x))/2 \) (where \( \Delta \) denotes the Laplacian) is lower bounded.

Assumption (1) is used to define a proposal distribution absolutely continuous with respect to the target which is easy to sample from. Assumption (2) is necessary to obtain a tractable Radon-Nikodym derivative between the proposal and the target distributions using the Girsanov transformation. While these assumptions can be proved under mild assumptions for scalar diffusions, much stronger conditions are required in the multidimensional case Ait-Sahalia (2008).

Let \( \omega = (\omega_s)_{0 \leq s \leq t} \) be the realization of a Brownian Bridge starting at \( x \) at time 0 and ending in \( y \) at time \( \Delta \). The distribution of \( \omega \) is denoted by \( \mathbb{W}_{\Delta}^x \). Moreover, suppose that for all \( \theta \in \Theta \), \( a_\theta \) is a gradient of a function \( A_\theta = \mathbb{V}^2 A_\theta \) where \( A_\theta : \mathbb{R}^d \to \mathbb{R} \) is a twice continuously differentiable function. Denoting, \( \psi_\theta : x \mapsto \psi_\theta(x) = (\|a_\theta(x)\|^2_+ + \Delta A_\theta(x))/2 \), by Girsanov theorem, for all \( x, y \in \mathbb{R}^d \times \mathbb{R}^d \)

\[
q_{k+1,\theta}(x, y) = \phi_{\Delta_k}(x - y) \exp \left( A_\theta(y) - A_\theta(x) \right) \mathbb{E}_{\mathbb{W}_{\Delta_k}^y} \left[ \exp \left( - \int_{0}^{\Delta_k} \psi_{\theta}(\omega_s) ds \right) \right],
\]

where \( \Delta_k = t_{k+1} - t_k \), for all \( a > 0 \), \( \phi_\theta \) is the probability density function of a centered Gaussian random variable with variance \( a \). The transition density then cannot be computed as it involves an integration over the whole path between \( x \) and \( y \). To perform the algorithm proposed in this article, we therefore have to design a positive unbiased estimator of \( q_{k+1,\theta}(x, y) \).

**Unbiased GPE estimator for \( q_{k+1,\theta}(x, y) \).** Assume that there exist random variables \( m_\theta \) and \( m_{\theta}\) such that for all \( 0 \leq s \leq \Delta_k \),

\[
m_\theta \leq \psi_\theta(\omega_s) \leq \overline{m}_\theta.
\]

Let \( \kappa \) be a random variable taking values in \( \mathbb{N} \) with distribution \( \mu = (\mu_k)_{0 \leq k \leq \Delta_k} \) be the realization of a Brownian Bridge, and \( (U_j)_{1 \leq j \leq \kappa} \) be independent uniform random variables on \((0, \Delta_k)\) and \( \xi = (x_0, y_0, U_1, \ldots, U_\kappa) \). As shown in Fearnhead et al. (2008), (21) leads to a positive unbiased estimator given by

\[
\hat{q}_{k+1,\theta}(x, y; \xi) = \phi_{\Delta_k}(x - y) \exp \left( A_\theta(y) - A_\theta(x) - m_\theta \Delta_k \right) \times \prod_{j=1}^{\kappa} \frac{\overline{m}_\theta - \psi_\theta(\omega_{U_j})}{m_\theta - m_{\theta}}.
\]

**Unbiased GPE estimator of \( \nabla_\theta \log q_{k+1,\theta}(x, y) \).** Let's denote \( \psi_\theta : x \mapsto \psi_\theta(x) - m_\theta \). By (21),

\[
\nabla_\theta \log q_{k+1,\theta}(x, y) = \nabla_\theta A_\theta(y) - \nabla_\theta A_\theta(x) - \nabla_\theta m_\theta \Delta_k - \mathbb{E}_{\mathbb{W}_{\Delta_k}^y} \left[ \int_{0}^{\Delta_k} \psi_{\theta}(\omega_s) ds \right] \times \prod_{j=1}^{\kappa} \frac{\overline{m}_\theta - \psi_{\theta}(\omega_{U_j})}{m_\theta - m_{\theta}}.
\]

On the other hand, the diffusion bridge \( \mathbb{W}_{\Delta_k}^y \) associated with the SDE (20) is absolutely continuous with respect to \( \mathbb{W}_{\Delta_k}^y \) with Radon-Nikodym derivative given by

\[
\frac{d\mathbb{W}_{\Delta_k}^y}{d\mathbb{W}_{\Delta_k}^x}(\omega) = \left[ q_{k+1,\theta}(x, y) \right]^{-1} \phi_{\Delta_k}(x - y) \exp \left( A_\theta(y) - A_\theta(x) - m_\theta \Delta_k - \int_{0}^{\Delta_k} \psi_{\theta}(\omega_s) ds \right) \times \mathbb{E}_{\mathbb{W}_{\Delta_k}^y} \left[ \exp \left( - \int_{0}^{\Delta_k} \psi_{\theta}(\omega_s) ds \right) \right]^{-1} \times \exp \left( - \int_{0}^{\Delta_k} \psi_{\theta}(\omega_s) ds \right).
\]

This yields

\[
\nabla_\theta \log q_{k+1,\theta}(x, y) = \left( \nabla_\theta A_\theta(y) - \nabla_\theta A_\theta(x) - \nabla_\theta m_\theta \Delta_k \right)
- \mathbb{E}_{\mathbb{W}_{\Delta_k}^y} \left[ \int_{0}^{\Delta_k} \nabla_\theta \psi_{\theta}(\omega_s) ds \right]
\times \exp \left( - \int_{0}^{\Delta_k} \psi_{\theta}(\omega_s) ds \right)
\]

and an unbiased estimator of \( \nabla_\theta \log q_{k+1,\theta}(x, y) \) is given by

\[
\hat{\nabla}_\theta \log q_{k+1,\theta}(x, y) = \left( \nabla_\theta A_\theta(y) - \nabla_\theta A_\theta(x) - \nabla_\theta m_\theta \Delta_k \right)
- \mathbb{E}_{\mathbb{W}_{\Delta_k}^y} \left[ \int_{0}^{\Delta_k} \nabla_\theta \psi_{\theta}(\omega_s) ds \right]
\times \exp \left( - \int_{0}^{\Delta_k} \psi_{\theta}(\omega_s) ds \right)
\]

where \( U \) is uniform on \((0, 1)\) and independent of \( \xi \). In the context of GPE, \( \hat{\nabla}_\theta \log q_{k+1,\theta}(x, y) \) can be computed exactly using exact algorithms for diffusion processes proposed in Beskos et al. (2006).

**A.3. Parametrix Estimators**

More recently, Andersson and Kohatsu-Higa (2017) and Fearnhead et al. (2017) proposed an algorithm which can be used under weaker assumptions. This parametrix algorithm draws weighted skeletons using an importance sampling mechanism for diffusion processes. In this case, the sampled paths are not distributed as the target process but the weighted samples produce unbiased estimates of expectations of functionals of this process. To obtain an unbiased estimator \( \hat{q}_{k+1,\theta}(x, y; \xi) \), the parametrix algorithm draws weighted skeletons at random times \( s_0 = 0 < s_1 < \ldots < s_j \), denoted by \( \left(x_{s_i}, w_{s_i}\right)\}_{j \geq 0} \) where \( x_0 = x \) and \( w_0 = 1 \). The update times \( s_{j+1} \) are instances of an inhomogeneous Poisson process of intensity \( \lambda(t) \).

Let \( x_{s_j}, w_{s_j} \) be the last weighted sample and \( s_{j+1} \) be the next update time of the trajectory. While \( s_{j+1} < \Delta_k \), the new state is sampled using a simple Euler scheme, namely:

\[
x_{s_{j+1}} := x_{s_j} + \Delta w_{s_j} \sigma_\theta(x_{s_j}) + (\Delta s_j)^{1/2} \sigma_\theta(x_{s_j}) w_{s_{j+1}},
\]

where \( \Delta s_j = s_{j+1} - s_j \), \( \Delta t_k = t_{k+1} - t_k \) and \( s_{j+1} \sim \mathcal{N}(0, I_d) \). The proposal density associated with this procedure is denoted by \( m_{\theta,\xi}(x_{s_j}, \Delta s_j) \). Let \( K_{\theta}^{\xi}(\xi) \) denote the Kolmogorov forward operator of the diffusion (resp. the Kolmogorov forward operator of the proposal distribution \( m_{\theta,\xi}(x_{s_j}, \Delta s_j) \)). The forward operators write, for any function \( h : \mathbb{R}^d \to \mathbb{R} \),

\[
K_{\theta}^{\xi} h(y) := - \sum_{i=1}^{d} \frac{\partial}{\partial y_i} \left[ a_{\theta,i}(y) h(y) \right] + \sum_{i,l=1}^{d} \frac{\partial^2}{\partial y_i \partial y_l} \left( \gamma_{\theta, i,l}(y) \nabla_{\theta,i} h(y) \right),
\]

where \( \gamma_{\theta, i,l} = \sigma_{\theta,i} \sigma_{\theta,l}^T \). Then, following Fearnhead et al. (2017), the weight is updated by

\[
w_{s_{j+1}} := w_{s_j} \rho_{\theta,\xi}^{s_{j+1}}(x_{s_j}, x_{s_{j+1}}, \Delta s_j),
\]
where

\[ \rho^\lambda_j(x,y,u) := 1 + \frac{\left(K - k_{\text{prop}}^j\right) m_{j\theta}(x,z,u)_{|z=y}}{\lambda(u) m_{j\theta}(x,y,u)}. \] (22)

It is worth noting that (22) can be computed using only first derivatives of \( \alpha_0 \) and second derivatives of \( \sigma_0 \). If \( N_k \) is the number of Poisson events between 0 and \( \Delta t_k \), the parametrix unbiased estimate is then given by

\[ \hat{q}_{k+1}(\xi_k)(x, y) = w_{N_k} m_k(\xi_k, \theta, t_{k+1} - s_{N_k}) \]

where \( \xi_k \) stands for all the randomness required to produce the parametrix estimator (Poisson process and Gaussian random variables).

The stability of this estimator is studied in Fearnhead et al. (2017) which provides \( L_p \) controls for the weight \( w_{N_k} \). The parametrix algorithm mentioned above is a highly flexible procedure to obtain such an unbiased estimate for a much broader class of diffusions than Poisson based estimations which require strong assumptions. However, as the update (22) involves the difference of two Kolmogorov operators, the parametrix estimator of the transition density may be negative, and has no reason to satisfy (13).

**Appendix B. Supplementary material for Section 6**

**B.1. Proposal Distribution in the Lotka-Volterra Experiment**

In this section, we detail the proposal distribution chosen for the particle filter applied to the Lotka-Volterra model defined in Section 6 of the article.

At each time \( k + 1 \), for \( k \geq 0 \), the objective is to sample a new particle \( \xi_{k+1} \) conditionally on a particle \( \xi_k \) at time \( k \), an observation \( Y_{k+1} \), and a parameter \( \theta \). In the following, \( \theta \) is dropped from the notations. Unlike in the Sine example, it is not straightforward to sample from the pdf proportional to the product

\[ w_{N_k} m_k(\xi_k, \theta, t_{k+1} - s_{N_k}) \phi_{\xi_k}(x_{k+1}), \]

as there is no explicit conjugation. A simple Gaussian approximation of the optimal filter is

\[ P_k(x_k, x_{k+1}) \propto m_k(x_k, x_{k+1}) \phi_{x_k}(x_{k+1}), \]

where \( m_k(\xi_k, y_{k+1}) \) the Gaussian pdf whose first moment is the one of \( Y_{k+1} | x_{k+1} \) given by (19) and variance \( \Sigma_k \) is chosen by the user. We denote \( m_{\text{filt}}(x_k, y_{k+1}) \) and \( \Sigma_{\text{filt}} \) the mean and variance of this approximate optimal Gaussian filter that can be easily computed. However, particles sampled from this distribution can take negative, which is not consistent with the model. As a workaround, \( \xi_k+1 \) is defined as \( \xi_k+1 \sim \exp(\xi_k)(x_{k+1}) \) where

\[ \xi_k+1 \sim \mathcal{N}(\ln m_{\text{filt}}(\xi_k, y_{k+1}) - \text{diag}(\Sigma_{\text{filt}})/2, \Sigma_{\text{filt}}). \]

The term \( \text{diag}(\Sigma_{\text{filt}})/2 \) in the mean of \( \xi_k+1 \) ensures that \( \mathbb{E} [\xi_k+1] = m_{\text{filt}}(\xi_k, y_{k+1}) \). In our experiments, \( \Sigma_{\text{prop}} \) was chosen as equal to \( \Sigma \).

**B.1.1. Estimation on Synthetic Data for the Lotka-Volterra Model**

Simulated data are obtained using the model given by (18) and (19) for a known set of parameters. Chosen values of \( \theta, \Sigma, c_1, \) and \( c_2 \) are the following:

\[ a_1 = (a_{11} = 12, a_{12} = 0.05, a_{13} = 1) \]

Prey dynamics parameters

\[ a_2 = (a_{21} = 2, a_{22} = 0.2, a_{23} = 0.1) \]

Predator dynamics param.

\[ \Gamma = (0.5, 0.1) \]

Diffusion parameters

\[ \mu_0 = (50, 20) \]

Mean of the initial state

\[ \Sigma_0 = I_2 \]

Variance of the initial state

\[ \Sigma_{\text{obs}} = (0.01, 0.005 \quad 0.005, 0.01) \]

Observation variance

\[ c = (0.2, 0.3) \]

Detectability values

The stochastic model is used to simulate abundances indexes \( Y_0, \ldots, Y_{300} \) at times \( t_0 = 0, \ldots, t_{300} = 3 \). The associated time series (after a division by the known constant \( c \)) is shown in Figure 8 (left panel). In this experiment, the goal is to obtain an estimate of the actual predator-prey abundances given all the observed abundances indexes \( Y_{0:n} \). Our estimate is given by the set of conditional expectations \( \mathbb{E}[X_k | Y_{0:n}] \) from a backward importance sampling PaRIS smoother, which is run using the true parameters. Figure 8 shows the estimated abundance trajectory over time. The proposed backward importance sampling smoother manages to estimate efficiently the actual abundance from noisy data and a model with an intractable transition density.

**B.2. Stochastic Volatility Model, Application to S & P 500 Data**

We consider a classical dataset of the daily log-returns of the Standard & Poor's 500 Index from 2005 to 2018, freely available in the R package stochvolTMB (Wahl 2021). Our goal is to estimate the volatility through a classical stochastic volatility model (Taylor 1982). Namely,
we consider the following model:

\[
X_k = \begin{cases} 
\sigma (1 - \phi^2)^{-1/2} \varepsilon_k & \text{for } k = 0, \\
\phi X_{k-1} + \sigma \varepsilon_k & \text{for } k > 0,
\end{cases}
\]

\[Y_k = \beta \exp \left( \frac{X_k}{2} \right) \zeta_k,
\]

where, $X_k$ is the hidden volatility at day $k$, $Y_k$ are the observed daily log-returns, $\beta, \phi$ and $\sigma$ are unknown positive parameters (and such that $\phi < 1$), and for $k \geq 0$, $(\varepsilon_k)_{k \geq 0}$ and $(\zeta_k)_{k \geq 0}$ are sequences of independent random variables with distribution $N(0, 1)$. We aim at performing maximum likelihood inference using an EM algorithm, where the $E$ step is approximated with our backward importance sampling algorithm. The $E$ step for this model relies on the approximation of five expectations of elementary additive functionals. Knowing these expectations, the $M$ step is then explicit (see, for instance, Douc et al. 2014, chap. 12, for computation details).

The particle smoother is implemented using $N = 400$ particles and $\tilde{N} = 200$.Estimated parameters are $\hat{\beta} = 0.01, \hat{\phi} = 0.98, \hat{\sigma} = 0.24$, which are consistent with usual estimations for this dataset (see for instance Wahl (2021)), and when performing the smoothing with the Forward Filtering Backward Simulation algorithm (Douc et al. 2011) or the Poor Man smoother. Figure 9 shows the resulting estimated volatility, where the exceptional periods such as the aftermath of the 2008 crisis are characterized by high volatility.

**Supplementary Materials**

Part of the code produced for the paper is available at: https://github.com/AMDonati/backward-IS.

**Disclosure Statement**

The authors report there are no competing interests to declare.

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