BAYESIAN RESTORATION OF HIGH-DIMENSIONAL PHOTON-STARVED IMAGES

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ABSTRACT

This paper investigates different algorithms to perform image restoration from single-photon measurements corrupted with Poisson noise. The restoration problem is formulated in a Bayesian framework and several state-of-the-art Monte Carlo samplers are considered to estimate the unknown image and quantify its uncertainty. The different samplers are compared through a series of experiments conducted with synthetic images. The results demonstrate the scaling properties of the proposed samplers as the dimensionality of the problem increases and the number of photons decreases. Moreover, our experiments show that for a certain photon budget (i.e., acquisition time of the imaging device), downsampling the observations can yield better reconstruction results.

Index Terms— Bayesian statistics, Inverse problems, Image processing, Poisson noise, Markov chain Monte Carlo, Bouncy particle sampler

1. INTRODUCTION

Photon-limited imaging arises in many applications where the light flux changes rapidly or is extremely limited. In such scenarios, the flux is quantified at photonic levels using single-photon detectors. A wide range of applications operate in this regime, such as light detection and ranging (lidar) [1,2], medical imaging [3] and astronomy [4]. Photon count measurements generally follow Poisson statistics, which vary significantly from the ones corrupted by Gaussian noise, yielding poor results if Gaussian restoration methods are directly applied. Moreover, algorithms based on the variance-stabilizing transformation fail to provide good estimates when the number of available photons is very low. In most photon-starved imaging applications, the number of recorded photons is proportional to the acquisition time. Thus, in order to reduce the acquisition time, the aim is to provide reliable restorations using as few photons as possible.

The problem can be stated as follows: Given a set of discrete photon measurements \( y = (y_1, ..., y_K)^T \in \mathbb{Z}_+^K = \{0,1,...\}^K \), we are interested in recovering the underlying intensities \( z = (z_1, ..., z_N)^T \in \mathbb{R}^N \) that are subjected to a linear operator \( A \in \mathbb{R}_+^{K \times N} \) such as blur, missing pixels, compressive measurements or tomographic projections, i.e.,

\[
y | z \sim P(Az).
\] (1)

This inverse problem (i.e., recovering \( z \) from \( y \)) is often severely ill-posed or ill-conditioned and prior regularization is necessary to promote the solution to be in a set of feasible images \( z \). Many convex optimization algorithms based on log-concave Bayesian models have been proposed to perform maximum-a-posteriori (MAP) estimation. For example, PI-DAL [5] is based on a total variation prior and solves the inverse problem using an alternating direction method of multipliers. Algorithms such as [5] are capable of computing MAP estimates relatively fast, but cannot provide uncertainty bounds for the estimates, which can be very valuable in many applications involving decision making such as self-driving cars [6]. In this paper, we formulate the inverse problem based on Eq. (1) in a Bayesian framework and investigate the efficiency of six different samplers at estimating the posterior distribution of \( z \). Specifically, we investigate four Markov chain Monte Carlo (MCMC) methods for the Poisson restoration problem: The unadjusted Langevin algorithm (ULA) [7], the Metropolis adjusted Langevin algorithm (MALA) [7], Hamiltonian Monte Carlo (HMC) [8, Chapter 5] and the recent no U-turn Hamiltonian Monte Carlo (NUTS) scheme [9]. In addition, we evaluate the bouncy particle sampler (BPS), a non-reversible rejection-free strategy [10]. For completeness, we also include in our comparison the classical random walk Metropolis (RWM) sampler [7]. Through a series of experiments, we study the scaling properties of each sampler as the dimension of the problem increases, i.e., as the size of the image increases and we evaluate the quality of the estimates as we vary the total number of photons. The main contribution of this paper is to investigate different stochastic simulation strategies in very high dimensional and very low photon count scenarios.

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2. BAYESIAN MODEL

Consider a vector of $K$ noisy measurements $y \in \mathbb{Z}_+^K$ and a vectorized intensity image $z \in \mathbb{R}^N$ of $n$ pixels. We are interested in the posterior distribution of the intensity values $z$ given the measurements $y$ denoted as $p(z|y)$. Following Bayes theorem, this posterior is given by

$$p(z|y) = \frac{p(y|z)p(z)}{\int p(y|z)p(z)dz}. \quad (2)$$

2.1. Poisson likelihood

The photon count of measurement $y_k$ is Poisson distributed

$$y_k|z \sim P\left(\sum_{n=1}^{N} A_{k,n} z_n\right) \quad k = 1, \ldots, K \quad (3)$$

where $P(.)$ is the Poisson distribution. Assuming independent noise realizations, the likelihood of $y$ is

$$p(y|z) = \prod_{k=1}^{K} p(y_k|z) \quad (4)$$

where $p(y_k|z)$ is the probability associated with the Poisson distribution of Eq. (3). To alleviate difficulties induced by the positivity constraint on $z$, we consider the reparametrization $z_n = e^{x_n}$ for $n = 1, \ldots, N$, where the log-intensity $x_n \in \mathbb{R}^N$ is no longer constrained to be positive, at the cost of inducing an implicit prior for $z$. The modified log-likelihood has a bounded Lipschitz constant as $x_n \rightarrow -\infty$, yielding better stability in low-photon count scenarios. A detailed study of the trade-off involved in this change of variables can be found in [11].

2.2. Laplacian filter prior

The target log-intensity vector $x = (x_1, \ldots, x_N)^T$ is regularized with the prior $p(x|\lambda) \sim N(0, (\lambda P)^{-1})$, where the precision matrix is $P = D^T D$ and $D$ is the block circulant matrix of the 2D Laplacian filter. This $l_2$ regularizer promotes smoothness in the log-image $x$. In this work, we choose this differentiable prior to allow for a wide variety of eligible samplers. In addition, it has been shown recently in [12] that this prior can be easily modified by changing the fixed Laplacian filter by other adaptive off-the-shelf denoisers to achieve state-of-the-art restoration performance. An in-depth study of this prior in a fully Bayesian configuration is out of the scope of this paper and is subject to future work.

2.3. Posterior distribution

Following Bayes rule, the posterior distribution of $x$ is

$$p(x|y, \lambda) \propto p(y|x)p(x|\lambda) \quad (5)$$

which can be written as $p(x|y, \lambda) = \frac{1}{C} \exp(-U_y(x))$, where $C$ is a normalizing constant and $U_y(x)$ is the negative log-posterior (or energy). Precisely,

$$U_y(x) = 1_T^T A e^x - y^T \log(A e^x) + \frac{\lambda}{2} x^T P x \quad (6)$$

where $[e^x]_n = e^{x_n}$ and $[\log(x)]_n = \log(x_n)$ are the element-wise exponential and logarithm vectors and $1_K = (1, \ldots, 1)^T$ is the $K \times 1$ vector of ones. Note that $U_y(x)$ is differentiable, convex and proper for any $y \in \mathbb{Z}_+^K$. If $A$ is injective, the energy $U_y(x)$ is also strictly convex with gradient

$$\nabla U_y(x) = e^x \odot \left(1_N - A^T (y^T \odot \frac{1}{A e^x}) \odot (A^T 1_K) + \lambda P x\right)$$

where $\frac{1}{A e^x}$ and $[x \odot y]_n = x_n y_n$ denote the element-wise vector inversion and multiplication.

3. ESTIMATORS AND INFERENCE STRATEGIES

The stochastic simulations methods presented in this work provide information of the full posterior distribution. In our experiments, we focus on two posterior statistics, the marginal posterior mean of each pixel defined as

$$E\{z_n|y, \lambda\} = \int_0^{+\infty} z_n p(z_n|y, \lambda) dz_n$$

which is the minimum mean squared error estimator (MMSE) of $z_n$ [13], and the marginal posterior variance

$$\text{var}\{z_n|y, \lambda\} = \int_0^{+\infty} (z_n - E\{z_n|y, \lambda\})^2 p(z_n|y, \lambda) dz_n. \quad (7)$$

These integrals are not analytically tractable. Thus we propose to compute numerical estimates $\hat{\mu}_n \approx E\{z_n|y, \lambda\}$ and $\hat{\sigma}_n^2 \approx \text{var}\{z_n|y, \lambda\}$, from samples of $p(x|y, \lambda)$ generated using MCMC samplers or particle trajectories (in the BPS case).

3.1. Markov chain Monte Carlo

MCMC samplers construct a Markov chain of $N_{MC}$ samples \{$x^{(s)}|s = 1, \ldots, N_{MC}$\}, distributed according to the posterior $p(x|y, \lambda)$. Discarding the first $N_{bi}$ burn-in iterations, the posterior statistics of interest are then computed as

$$\hat{\mu}_n = \frac{1}{N_{MC}} \sum_{s=N_{bi}}^{N_{MC}} x_n^{(s)} \quad (7)$$

and

$$\hat{\sigma}_n^2 = \frac{1}{N_{MC}} \sum_{s=N_{bi}}^{N_{MC}} [2x_n^{(s)} - \hat{\mu}_n^2]. \quad (8)$$

We consider five different MCMC samplers: the classical RWM with a data-independent proposal and four gradient-based samplers, ULA, MALA, HMC and NUTS. In the case
of RWM and MALA, we adjust the parameter controlling
the variance of the proposal distribution during the burn-in
to achieve an acceptance rate close to the optimal one, as
explained in [8, Chapter 4]. We used the adaptive scheme
described in [14], in order to obtain an acceptance rate close
to 0.23 for RWM and 0.57 for MALA. For HMC, we fix the
path length and adapt the step size using the algorithm
described in [9] to obtain an acceptance rate close to the optimal
one of 0.65. NUTS controls both step size and number of
steps adaptively to yield acceptance rates close to 0.65.

3.2. Bouncy Particle Sampler

A non-reversible rejection-free sampler was recently pro-
posed in [10], performing better than the optimally tuned
HMC for some Bayesian models. This algorithm simulates a
particle that travels around the posterior with constant veloc-
ity $v \in \mathbb{R}^N$ and performs elastic bounces with a hyperplane
of normal vector $\nabla U_{\text{y}}(x)$. If the energy is strictly convex, the
$s$th bouncing time $\tau^s$ is computed by solving the one
dimensional equation

$$U_{\text{y}}(x^s) + \nu^s \tau^s - U_{\text{y}}(x^s) + \nu^s \tau^s = -\log V \quad (9)$$

where $\tau^s = \arg \min_{t \geq 0} U_{\text{y}}(x^s) + \nu^s t \quad (10)$

and $V \sim U(0, 1)$. In our configuration, the solutions of Eq.
(9) and (10) cannot be computed analytically and we use two
one-dimensional Newton iterative methods to find them. The
particle trajectory is ergodic, allowing estimates of $\mu_n$ and $\sigma_n^2$
to be computed as time-averages of the trajectory, i.e.,

$$\hat{\mu}_n = \frac{1}{N_b} \sum_{s=1}^{N_b} \exp \left[ \frac{U_{\text{y}}(x^s) + \nu^s \tau^s}{\nu_n} \right]$$

(11)

and

$$\hat{\sigma}_n^2 = \frac{1}{N_b} \sum_{s=1}^{N_b} \exp \left[ \frac{U_{\text{y}}(x^s) + 2 \nu^s \tau^s}{2 \nu_n} \right] - \hat{\mu}_n^2 \quad (12)$$

where $N_b$ is the total number of simulated bounces. To as-
sure convergence, the particle also performs random bounces
according to a user-specified refreshment rate, which was set
to 1 by cross-validation (see [10] for additional details about
the BPS algorithm).

3.3. Computational complexity

In a general setting, the matrix $A$ can be dense (e.g., built
from many tomographic projections or degraded by a large
blurring kernel). The precision matrix $P$ can also be dense, as
the prior can account for non-local correlations between im-
age patches [15]. Although this is not the case for the Lapla-
cian filter, we will keep the dense assumption for the sake of
generality. Under such assumptions, the conditional distribu-
tion of any pixel $z_n$ depends on many neighbours $z_i$, $i \neq n$.
Thus, techniques that exploit a sparse posterior factor graph
(e.g., Gibbs sampling using a checkerboard scheme [16], lo-
cal bouncy particle sampler [10], Hamiltonian splitting ap-
proaches [8, Chapter 5]) are not applicable. For each sample
generated by an MCMC method or each bounce for NUTS,
the algorithm has to compute $U(x)$ and/or $\nabla U(x)$, which are
dominated by the cost of evaluating the forward operator
$A e^x$ and the Laplacian prior term $P x$. The forward oper-
ator $A$ requires $N_A$ flops of order $O(NK)$, which reduces
to $O(N \log N)$ if $A$ is block-circulant and the Fast Fourier
transform (FFT) algorithm is used. A similar cost is related
to the prior term $P x$, as it involves one filtering operation
of $N_P$ flops with order $O(N \log N)$. Note that the rest of
the operations have order $O(N)$. We are interested in obtain-
ing good estimators for a small number of $N_P$ and $N_A$ flops.
Table 3.3 summarizes the number of per-sample (or bounce)
computations of each sampler.

<table>
<thead>
<tr>
<th>Sampler</th>
<th>Complexity</th>
</tr>
</thead>
<tbody>
<tr>
<td>RWM</td>
<td>$N_A + N_P$</td>
</tr>
<tr>
<td>ULA</td>
<td>$2N_A + N_P$</td>
</tr>
<tr>
<td>MALA</td>
<td>$2N_A + N_P$</td>
</tr>
<tr>
<td>HMC</td>
<td>$L(2N_A + N_P)$</td>
</tr>
<tr>
<td>NUTS</td>
<td>$L(2N_A + N_P)$</td>
</tr>
<tr>
<td>BPS</td>
<td>$(3k_1 + 2k_2)N_A + N_P$</td>
</tr>
</tbody>
</table>

Table 1. Complexity per iteration of the proposed samplers ($k_1, k_2$ are the numbers of Newton steps for Eq. (9) and (10)).

4. EXPERIMENTS

We investigated the performance of the samplers as functions
of the total number of photon counts and the dimension-
ality of the problem. All the experiments were performed with
the “cameraman” image. We fixed the parameter controlling
the amount of regularization $\lambda = 2$ by cross-validation.

Figure 1 shows the true intensity image and a noisy realiza-
tion with $10^4$ photons. The burn-in was set to 30% of the
computing time in all the simulations. First, we compared

Fig. 1. Left: True intensity image. Right: Noisy observations
for a $256 \times 256$ image and $10^4$ photons.
the bias and variance of the estimates $\hat{\mu}_n$ and $\hat{\sigma}^2_n$ of a randomly chosen pixel for image sizes of $64 \times 64$, $128 \times 128$ and $256 \times 256$ pixels with a fixed photon budget of $10^4$ photons and an operator $A$ defined as a blur of $3 \times 3$ pixels. Each sampler was run 10 times using different random seeds. Figures 2 and 3 show the means and standard deviations of the estimates $\hat{\mu}_n$ and $\hat{\sigma}^2_n$ respectively. HMC achieves the best performance overall, converging without bias in less computing time. NUTS performs slightly worse, as the number of steps per iteration increases significantly faster than HMC with the image size, thus leading to less samples and worse estimates. MALA and ULA have reasonable performance, but ULA shows biased estimates as it does not converge to the exact posterior distribution [17]. The estimates of the means converge relatively fast for BPS, but the estimates of the variances show markedly slower convergence with respect to the other MCMC gradient-based samplers. RWM has the worst performance for this example.

Second, we fixed the image size to $64 \times 64$ and compared the samplers versus the number of photon counts. Figure 4 shows the results for photon counts of $10^5$ and $10^3$. When the photon count is lower, the likelihood term Eq. (3) flattens, meaning longer posterior tails and more samples needed to achieve good estimation performance. Again, HMC provides the best results for a given computing time.

The sum of Poisson random variables is Poisson dis-
tributed. Hence, binning the noisy image with a $2 \times 2$ pixels window yields a 4 times increase in signal-to-noise ratio, at the cost of lower high frequency information. Recent work in the Poisson denoising literature has shown that lower scales can add valuable information [18, 19]. We compared the normalized mean squared error

$$\text{NMSE} = \frac{\sum_{n=1}^{N} (z_n - \hat{\mu}_n)^2}{\sum_{n=1}^{N} \mu_n^2}$$

obtained by the HMC sampler for different downsampled scales. In this experiment, we removed the forward operator (i.e., the matrix $A$ was set to the identity matrix) and fixed the computing time to 300 seconds. The coarse scale results were upsampled by simply assigning 1/4 of the estimated intensity to each pixel in the finer scale. Table 2 shows the NMSE of each scale from $10^2$ photons to $5 \times 10^5$ photons. In a very low-photon count scenario (less than $10^4$ photons), the estimates at lower scales perform better than at higher ones, also requiring less time to converge. When the photon number increases, the results are better at finer scales, as they contain more high-frequency information.

<table>
<thead>
<tr>
<th>photons/ $p$</th>
<th>16</th>
<th>32</th>
<th>64</th>
<th>128</th>
<th>256</th>
</tr>
</thead>
<tbody>
<tr>
<td>$10^2$</td>
<td>14.1</td>
<td>16.4</td>
<td>29.7</td>
<td>67.62</td>
<td>200.7</td>
</tr>
<tr>
<td>$10^3$</td>
<td>10.9</td>
<td>10.7</td>
<td>12.08</td>
<td>16.48</td>
<td>47</td>
</tr>
<tr>
<td>$10^4$</td>
<td>7.5</td>
<td>6.7</td>
<td>6.5</td>
<td>7.3</td>
<td>10.7</td>
</tr>
<tr>
<td>$2 \times 10^4$</td>
<td>6.9</td>
<td>5.9</td>
<td>5.4</td>
<td>5.8</td>
<td>7.9</td>
</tr>
<tr>
<td>$5 \times 10^4$</td>
<td>6.5</td>
<td>5.2</td>
<td>4.5</td>
<td>4.4</td>
<td>6.5</td>
</tr>
<tr>
<td>$7 \times 10^4$</td>
<td>6.4</td>
<td>5.0</td>
<td>4.1</td>
<td>3.9</td>
<td>4.5</td>
</tr>
<tr>
<td>$5 \times 10^5$</td>
<td>6.0</td>
<td>4.3</td>
<td>2.9</td>
<td>2.2</td>
<td>2.0</td>
</tr>
</tbody>
</table>

Table 2. NMSEs at different scales (the image size is $p \times p$).

5. CONCLUSION

This paper studied a Bayesian approach for image restoration using low-photon count measurements. Six different stochastic simulation algorithms were compared to compute posterior statistics of interest, in particular the marginal pixel mean and variance, for this image restoration problem. The experiments performed with synthetic data showed that the Hamiltonian Monte Carlo sampler provides better estimates as the image size increases and the number of photon counts decreases. These results are in agreement with the theoretical scaling of the effective sample size [8] [10]. Finally, we showed experimentally the trade-off between the available photon budget and the scale associated with the best restoration. When the number of photon counts is very low compared to the image size, downsampling the image can yield better results using less computing time.

6. REFERENCES


