The Wang-Landau algorithm in general state spaces: applications and convergence analysis

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THE WANG-LANDAU ALGORITHM
IN GENERAL STATE SPACES:
APPLICATIONS AND CONVERGENCE ANALYSIS

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University of Michigan and Harvard University

Abstract: The Wang-Landau algorithm (Wang and Landau (2001)) is a recent Monte Carlo method that has generated much interest in the Physics literature due to some spectacular simulation performances. The objective of this paper is two-fold. First, we show that the algorithm can be naturally extended to more general state spaces and used to improve on Markov Chain Monte Carlo schemes of more interest in Statistics. In a second part, we study asymptotic behaviors of the algorithm. We show that with an appropriate choice of the step-size, the algorithm is consistent and a strong law of large numbers holds under some fairly mild conditions. We have also shown by simulations the potential advantage of the WL algorithm for problems in Bayesian inference.

Key words and phrases: Adaptive MCMC, geometric ergodicity, Monte Carlo methods, multicanonical sampling, stochastic approximation, trans-dimensional MCMC, Wang-Landau algorithm.

1. Introduction

Although the idea of Monte Carlo computation has been around for more than a century, its first real scientific use occurred during the World War II when the first generation computer became available. Nick Metropolis coined the name “Monte Carlo” for the method when he was at Los Alamos National Labs and it quick evolved to an active research area due to the active involvements of leading physicists in the Labs. Ever since then, physicists have been at the forefront of the methodological research in the field. One of their latest additions is an algorithm proposed by F. Wang and D. P. Landau (Wang and Landau (2001)). The Wang-Landau (WL) algorithm has been successfully applied to some complex sampling problems in physics. The algorithm is closely related to multicanonical sampling, a method due to B. A. Berg and T. Neuhaus (Berg and Neuhaus (1992)). Briefly, if $\pi$ is the probability measure of interest, the idea behind multicanonical sampling is to obtain an importance sampling distribution by partitioning the state space along the energy function ($-\log \pi(x)$) and re-weighting
appropriately each component of the partition so that the modified distribution \( \pi^* \) spends equal amount of time in each component, i.e., uniform in the energy space. The method is often criticized for the difficulty involved in computing the weights. The main contribution of the WL algorithm is in proposing an efficient algorithm that simultaneously computes the balancing weights and samples from the re-weighted distribution.

The objective of this paper is to take a more probabilistic look at the WL algorithm and to explore its potential for Monte Carlo simulation problems of more direct interest to statisticians. We achieve this goal by proposing a general state space version of the algorithm. Then we show that the WL algorithm offers an effective strategy to improve on simulated tempering and trans-dimensional MCMC.

From a probabilistic standpoint, the WL algorithm is an interesting example of adaptive Markov Chain Monte Carlo (MCMC). Adaptive MCMC is an approach to Monte Carlo simulation where the transition kernel of the algorithm is sequentially adjusted over time in order to achieve some prescribed optimality. Some early work on the subject includes [Gilks, Roberts and Sahu (1998), Andrieu and Robert (2001), and Haario, Saksman and Tamminen (2001)]. See also [Brockwell and Kadane (2005), and Mira and Sargent (2003)]. Early theoretical analysis includes [Atchadé and Rosenthal (2005), Andrieu and Moulines (2006), Andrieu and Atchade (2007), and Rosenthal and Roberts (2007)]. We take a similar path-wise approach to analyze the WL algorithm. The analysis of the WL algorithm is not a straightforward application of the theory in the aforementioned papers because of the specific adaptive control involved. The key point is the stability of the algorithm. We say that the WL algorithm is stable if no component of the partition receives infinitely more visits than any other component as \( n \to \infty \). On a stable sample path, and under appropriate conditions, we show that the WL algorithm learns the optimal weights and satisfies a strong law of large numbers (Theorem 4.1). In the specific cases of multicanonical sampling and simulated tempering, which includes the original WL algorithm, we show that the algorithm is stable and that the aforementioned limit results hold.

It came to our attention after the first draft of this paper that a similar extension of the WL algorithm has been proposed independently by F. Liang and coworkers (see e.g., Liang, Liu and Carroll (2007)). Their approach differs from the WL approach in that these authors took a more classic approach based on stochastic approximation with step-sizes set deterministically.

Our proposed generalization to the WL algorithm is presented in Section 2. Some particular cases are discussed in Section 3. The theoretical analysis is discussed in Section 4, but the proofs are postponed to Section 6 to facilitate the flow of ideas.
2. The Wang-Landau Algorithm

In multicanonical sampling, we are given a state space $\mathcal{X}$ and a probability measure $\pi$. $\mathcal{X}$ is then partitioned as $\mathcal{X} = \bigcup_{i=1}^{d} \mathcal{X}_i$, where $\mathcal{X}_i \cap \mathcal{X}_j = \emptyset$ and $\pi$ is re-weighted in each component $\mathcal{X}_i$. An abstract way to do the same and much more is the following. We start with $(\mathcal{X}_i, \mathcal{B}_i, \lambda_i)_{i=1}^{d}$, a finite family of measure spaces where $\lambda_i$ is a $\sigma$-finite measure. We introduce the union space $\mathcal{X} = \bigcup_{i=1}^{d} \mathcal{X}_i \times \{i\}$. We equip $\mathcal{X}$ with the $\sigma$-algebra $\mathcal{B}$ generated by $(A_i, i, A_i \in \mathcal{B}_i)$ and the measure $\lambda$ satisfying $\lambda(A, i) = \lambda_i(A)1_{\mathcal{B}_i}(A)$. Let $h_i : \mathcal{X}_i \rightarrow \mathbb{R}$ be a non-negative measurable function and define $\theta^*(i) = \int_{\mathcal{X}_i} h_i(x) \lambda_i(dx)/Z$ where $Z = \sum_{i=1}^{d} \int_{\mathcal{X}_i} h_i(x) \lambda_i(dx)$. We assume that $\theta^*(i) > 0$ for all $i = 1, \ldots, d$, and consider the following probability measure on $(\mathcal{X}, \mathcal{B})$:

$$
\pi^*(dx, i) \propto \frac{h_i(x)}{\theta^*(i)} 1_{\mathcal{X}_i}(x) \lambda_i(dx).
$$

Our objective is to sample from $\pi^*$. The problem of sampling from such a distribution arises in a number of different Monte Carlo strategies. For example, and as explained above, if $\pi$ is a probability measure of interest on some space $(\mathcal{X}, \mathcal{B}, \lambda)$, we can partition $\mathcal{X}$ along the energy function $-\log(\pi)$ and re-weight $\pi$ by $\pi(\mathcal{X}_i)$ in each component $\mathcal{X}_i$. The sampling problem then becomes of the form (2.1). This powerful strategy appeared first in the Physics literature as multicanonical sampling (Berg and Neuhaus (1992)). This is discussed in some more details in Section 3.1.

Sampling from (2.1) also arises naturally when optimizing the simulated tempering algorithm (Marinari and Parisi (1992), and Geyer and Thompson (1995)). In simulated tempering, the states space $\mathcal{X}$ is not partitioned but, instead, some auxiliary distributions $\pi_2, \ldots, \pi_d$ are introduced (take $\pi_1 = \pi$). These distributions are chosen close to $\pi$ but easier to sample from. For good performances, one typically imposes that all the distributions have the same weight. Taking each probability space $(\mathcal{X}, \mathcal{B}, \pi_i)$ as a component in the formalism above leads to a sampling problem of the form (2.1). Multicanonical sampling and simulated tempering have been combined in Atchadé and Liu (2006) giving an algorithm which can also be framed as (2.1). Sampling from (2.1) can also be an efficient strategy to improve on trans-dimensional MCMC samplers for Bayesian inference with model uncertainty. This is detailed in Section 3.3.

The main obstacle in sampling from $\pi^*$ is that the normalizing constants $\theta^*$ are not known. The contribution of the Wang-Landau algorithm (Wang and Landau (2001)) is an efficient algorithm that simultaneously estimates $\theta^*$ and sample from $\pi^*$. The algorithm was introduced in a discrete setting with the $\pi^*$
being uniform in $i$. In this work we extend the algorithm to general state spaces and to arbitrary probability measures. To carry on the discussion in our general framework, we introduce the family of probability measures $\{\pi_\theta, \theta \in (0, \infty)^d\}$ on $(\mathcal{X}, \mathcal{B}, \lambda)$ defined by:

$$
\pi_\theta(dx, i) \propto \frac{h_i(x)}{\theta(i)} 1_{\mathcal{X}_i}(x) \lambda(dx).
$$

We assume that for all $\theta \in (0, \infty)^d$, we have at our disposal a transition kernel $P_\theta$ on $(\mathcal{X}, \mathcal{B})$ with invariant distribution $\pi_\theta$. Note that $\pi_\theta$ and $P_\theta$ remain unchanged if we multiply the vector $\theta$ by a positive constant. How to build such Markov chain $P_\theta$ typically depends on the particular instance of the algorithm. We give some examples later.

The structure of the WL algorithm is as follows. We start out with some initial value $(X_0, I_0) \in \mathcal{X}$, $\phi_0 \in (0, \infty)^d$, and set $\theta_0(i) = \phi_0(i)/\sum_j \phi_0(j)$, $i = 1, \ldots, d$. Here $\theta_0$ serves as an initial guess of $\theta^*$. At iteration $n + 1$, we generate $(X_{n+1}, I_{n+1})$ by sampling from $P_\phi_n(X_n, I_n; \cdot)$ and update $\phi_n$ to $\phi_{n+1}$, which is used to form $\theta_{n+1}(i) = \phi_{n+1}(i)/\sum_j \phi_{n+1}(j)$. The updating rule for $\phi_n$ is fairly simple. For $i \in \{1, \ldots, d\}$, if $X_{n+1} \in X_i$ (equivalently, if $I_{n+1} = i$), then $\phi_{n+1}(i) = \phi_n(i)(1 + \rho)$ for some $\rho > 0$; otherwise $\phi_{n+1}(i) = \phi_n(i)$. This leads to a first version of the WL algorithm.

**Algorithm 2.1.** (The Wang-Landau algorithm I). Let $\{\rho_n\}$ be a sequence of decreasing positive numbers. Let $(X_0, I_0) \in \mathcal{X}$ be given. Let $\phi_0 \in \mathbb{R}^d$ be such that $\phi_0(i) > 0$, and set $\theta_0(i) = \phi_0(i)/\sum_j \phi_0(j)$, $i = 1, \ldots, d$. At some time $n \geq 0$, given $(X_n, I_n) \in \mathcal{X}$, $\phi_n \in \mathbb{R}^d$, $\theta_n \in \mathbb{R}^d$:

(i) sample $(X_{n+1}, I_{n+1}) \sim P_\phi_n(X_n, I_n; \cdot)$;

(ii) for $i = 1, \ldots, d$, set $\phi_{n+1}(i) = \phi_n(i)(1 + \rho_n 1_{\{I_{n+1} = i\}})$ and $\theta_{n+1}(i) = \phi_{n+1}(i)/\sum_j \phi_{n+1}(j)$.

It remains to choose the sequence $\{\rho_n\}$. As we show below, $\{\theta_n\}$ as defined by Algorithm 2.1 is a stochastic approximation process driven by $\{(X_n, I_n)\}$. The general guidelines in the literature to choose $\{\rho_n\}$ are: $\rho_n > 0$, $\sum \rho_n = \infty$, and $\sum \rho_n^{1+\varepsilon} < \infty$ for some $\varepsilon > 0$, often $\varepsilon = 1$. The typical choice is $\rho_n \propto n^{-1}$. In practice, more careful choices are often necessary for good performance. To the best of our knowledge, there is no general, satisfactory way of choosing the step-size in stochastic approximation. Interestingly, Wang-Landau came up with a clever, adaptive way of choosing $\{\rho_n\}$ that works very well in practice. We describe their approach next, again in more probabilistic terms.

Let $v_{n,k}(i)$ denote the proportion of visits to $\mathcal{X}_i \times \{i\}$ between times $n + 1$ and $k$. That is, $v_{k,n}(i) = 0$ for $k \leq n$ and for $k \geq n + 1$, $v_{n,k}(i) = 1/(k -
At some time $n \mid \sum_{j=n+1}^{n} 1 \{I_j = i\}$. Let $c \in (0, 1)$ be a parameter to be specified by the user. We introduce two additional random sequences $\{\kappa_n\}$ and $\{a_n\}$. Initially, $\kappa_0 = 0$. For $n \geq 1$, define

$$\kappa_n = \inf \left\{ k > \kappa_{n-1} : \max_{1 \leq i \leq d} \left| v_{\kappa_{n-1},k}(i) - \frac{1}{d} \right| \leq \frac{c}{d} \right\}, \quad (2.3)$$

with the usual convention that $\inf \emptyset = \infty$. We need another sequence $\{\gamma_n\}$ of positive decreasing numbers, representing “stepsizes”. Then, $\{a_n\}$ represents the index of the element of the sequence $\{\gamma_n\}$ used at time $n : a_0 = 0$, if $k = \kappa_j$ for some $j \geq 1$, then $a_k = a_{k-1} + 1$, otherwise $a_k = a_{k-1}$. In words, we start Algorithm 2.1 with a step-size equal to $\gamma_0$ and keep using it until time $\kappa_1$ when all the components are visited equally well. Only then do we change the step-size to $\gamma_1$ and keep it constant until time $\kappa_2$ etc... Combining this with Algorithm 2.1, we get the following.

**Algorithm 2.2.** (The Wang-Landau algorithm II). Let $\{\gamma_n\}$ be a sequence of decreasing positive numbers. Let $(X_0, I_0) \in \mathcal{X}$ be given. Set $a_0 = 0$, $\kappa = 0$, $c \in (0, 1)$, $\phi_0 \in \mathbb{R}^d$ such that $\phi_0(i) > 0$, and $\theta_0(i) = \phi_0(i)/\sum_j \phi_0(j)$, $i = 1, \ldots, d$. At some time $n \geq 0$, given $(X_n, I_n) \in \mathcal{X}$, $\phi_n \in \mathbb{R}^d$, $\theta_n \in \mathbb{R}^d$, $a_n$ and $\kappa$:

(i) sample $(X_{n+1}, I_{n+1}) \sim P_{\theta_n}(X_n, I_n; \cdot)$;

(ii) for $i = 1, \ldots, d$, set $\phi_{n+1}(i) = \phi_n(i)(1 + \gamma_{a_n} 1_{\{I_{n+1} = i\}}) \quad \text{and} \quad \theta_{n+1}(i) = \phi_{n+1}(i) / \sum_j \phi_{n+1}(j)$;

(iii) if $\max_i |v_{\kappa_{n+1},1}(i) - (1/d)| \leq c/d$ then set $\kappa = n + 1$ and $a_{n+1} = a_n + 1$, otherwise $a_{n+1} = a_n$.

**Remark 2.1.**

1. The performances of Algorithm 2.2 depends very much on the choice of $\{\gamma_n\}$ and $c$. We show that the choice $\gamma_n \propto n^{-1}$ guaranties the convergence of the algorithm but, in practice, this type of step-size can be overly slow. The user might then consider $\gamma_n \propto a^{-n}$ $(a > 1)$ originally proposed by Wang-Landau, but we were not able to obtain the convergence of the algorithm for summable step-sizes. A good compromise is to start the sampler with $\gamma_n \propto a^{-n}$ until $\gamma_n < \varepsilon$ (e.g., $\varepsilon = 10^{-5}$) and then switch to $\gamma_n = n^{-1}$. There is a bias-variance trade-off involved in the choice of $c$. For $c$ close to 0, Algorithm 2.2 will have a low bias (in estimating $\theta^*$) but a high variance. Such values of $c$ are more suitable for $\gamma_n \propto a^{-n}$ whereas for larger values of $c$, the bias will be high with a low variance. For $c = d - 1$, we get a standard stochastic approximation algorithm with deterministic step-size for which a step-size that sums to $\infty$ is necessary for convergence. We found empirically from our simulations that $c$ in the range $0.4 - 0.2$ yields reasonably good samplers.
In the actual implementation of the algorithm, it is not necessary to re-normalize \( \phi_n \) into \( \theta_n \) as in (ii). In fact, for computational stability, we recommend carrying out the recursion on a logarithmic scale: 
\[
\log \phi_n(i) = \log \phi_{n-1}(i) + \log(1 + \gamma_{n-1}) \mathbf{1}_{\{I_n=i\}}.
\]

Another interesting feature of Algorithm 2.2 is that Step (iii) can serve as a stopping rule: we stop the simulation when \( \gamma_n \) get smaller than some pre-specified value.

Under some regularity conditions, if \( f : \mathcal{X} \to \mathbb{R} \) is some function of interest that is \( \pi \)-integrable and \( \{(X_n, I_n, \theta_n)\} \) is as described in Algorithm 2.2, we will show below that
\[
\frac{1}{n} \sum_{k=1}^{n} f(X_k, I_k) \to \pi^*(f), \operatorname{a.s.} \text{ as } n \to \infty.
\]

If we denote by \( \pi_i \) the distribution on \( \mathcal{X}_i \) with density with respect to \( \lambda_i \) proportional to \( h_i(x) \mathbf{1}_{\mathcal{X}_i}(x) \), we can estimate integrals with respect to \( \pi_i \) as well:
\[
\frac{\sum_{k=1}^{n} f(X_k, I_k) \mathbf{1}_{\mathcal{X}_i}(X_k)}{\sum_{k=1}^{n} \mathbf{1}_{\mathcal{X}_i}(X_k)} \to \pi_i(f(\cdot, i)), \operatorname{a.s.} \text{ as } n \to \infty.
\]

Now if we denote by \( \pi \) the distribution on \( \mathcal{X} \) whose density with respect to \( \lambda \) is proportional to \( h_i(x) \) on \( \mathcal{X}_i \), the ratio \( \pi(x, i) / \pi^*(x, i) \) is \( d\theta^*(i) \) and integrals with respect to \( \pi \) can also computed by importance sampling:
\[
\frac{d}{n} \sum_{k=1}^{n} f(X_k, I_k) \theta_k(I_k) \to \pi(f), \operatorname{a.s.} \text{ as } n \to \infty.
\]

Various methods for recycling Monte Carlo samples can be implemented as well. All these results follow from Theorem 4.1.

### 3. Some Applications

In this section, we detail briefly some applications of the general algorithm to multicanonical sampling, simulated tempering, and trans-dimensional MCMC.

#### 3.1. Multicanonical sampling

Multicanonical sampling is a powerful algorithm proposed by Berg and Neuhaus (1992). It holds the potential of improving on mixing times of classical MCMC algorithms. It fits naturally in the framework above, but the implementation can be tedious. Assume that we want to sample from a probability measure \( \pi(dx) \propto h(x)\lambda(dx) \) on some probability space \((\Sigma, \mathcal{A}, \lambda)\). We use the energy function \( E(x) = -\log(h(x)) \) to build a \( d \)-component partition \( \{\mathcal{X}_i\}_i \) of \( \Sigma \), \( \mathcal{X}_i = \{x \in \Sigma : E_{i-1} < E(x) \leq E_i\} \), where \( -\infty \leq E_0 < E_1 < \cdots < E_d \leq \infty \).
are predefined values. Let $\theta^*(i) = \pi(X_i)$ and assume $\theta^*(i) > 0$. As above, we introduce the union space $X = \bigcup X_i \times \{i\}$. The idea of multicanonical sampling is to sample from

$$\pi^*(dx, i) \propto \frac{h(x)}{\theta^*(i)} 1_{X_i}(x) \lambda(dx),$$

which is of the form (2.1). There is a simpler formulation of the algorithm. Since the component of the partition to which a point $x$ belongs can be obtained from $x$ itself, multicanonical sampling is equivalent to sampling from $\pi$ on $(\Sigma, A)$ given by

$$\pi^*(dx) \propto \sum_{i=1}^d \frac{h(x)}{\theta^*(i)} 1_{X_i}(x) \lambda(dx),$$

(3.1)

and the union space formalism is not needed. After sampling from $\pi^*$ in (3.1), a straightforward importance sampling estimate allows one to recover $\pi$. The algorithm tries to break the barriers in the energy landscape of the distribution by re-weighting each component $X_i$. Clearly, the success depends heavily on a good choice of the energy rings $E_0, \ldots, E_d$. This typically requires some prior information on $\pi$ or some pilot simulations. We point out that, although the energy function $E$ is a natural candidate to utilize to partition the space, the idea can be extended to other functions.

In the description of multicanonical sampling given above, taking $\Sigma$ as a discrete space, $\pi$ the uniform distribution on $X$, and $X_e = \{x \in \Sigma : E(x) = e\}, e \in \{e \in \mathbb{R} : E(x) = e \text{ for some } x \in \Sigma\}$ yields the Wang-Landau algorithm of [Wang and Landau (2001)].

### 3.2. Simulated tempering

The method can be applied to the simulated tempering of [Marinari and Parisi (1992)] and [Geyer and Thompson (1995)] by taking $(X_i, B_i, \lambda_i) \equiv (X_1, B_1, \lambda_1)$ and $h_i = h^{1/t_i}, \ 1 = t_1 < \cdots < t_d$. Simulated tempering is a well-known Monte Carlo strategy for sampling from difficult target distributions. Assume that the distribution of interest is $\pi_1(dx) \propto h(x) \lambda(dx)$. Typically for large temperature $t$, $h^{1/t}$ is a more well-behaved distribution for which faster mixing Markov chains can be built. In simulated tempering, we try to take advantage of these faster mixing chains by targeting the distribution

$$\pi_\theta(dx, i) = \frac{h^{1/t_i}(x)}{\theta^*(i)} 1_{X_i}(x) \lambda_1(dx),$$

(3.2)

on the union space $(X, B, \lambda)$. A MCMC sampler $P_\theta$ with invariant distribution $\pi_\theta$ is readily designed. Typically $P_\theta$ takes the form

$$P_\theta((x, i); A \times \{j\}) = B_{x, \theta}(i, j) P^{[j]}(x, A),$$

(3.3)
where $B_{x,\theta}$ is a transition kernel on $\{1, \ldots, d\}$ with invariant distribution 
\[ (h_i(x)/\theta(i))(\sum_i h_i(x)/\theta(i))^{-1} \] and $P^{[i]}$ a transition kernel on $(X_1, B_1)$ with

invariant distribution proportional to $h_i(x)\lambda(dx)$. Typically, one takes $B_{x,\theta}(i, j) = (h_i(x)/\theta(i))(\sum_i h_i(x)/\theta(i))^{-1}$. Another common choice is to take $B_{x,\theta}$ as a

Metropolis-Kernel on $f_1, \ldots, d$ with proposal $q(i, j)$. By standard importance

sampling techniques, we can convert samples from the higher temperature dis-

tributions to estimate $\pi_1$. The method holds for any $\theta \in \mathbb{R}_d$, $\theta(i) > 0$, but the

choice of $\theta$ can significantly impact the efficiency. Heuristically it seems that, to

improve on mixing, we need a $\theta$ that allows samples from fast converging dis-

tributions (but close to $\pi$); but since $\pi_1$ is the distribution of interest, for statistical

efficiency, we need a $\theta$ that favors $\pi_1$. One easy way to resolve this trade-off is

to choose $\theta(i) = \theta^*(i) \propto \int h_i(x)\lambda_1(dx)$ and sample from

$\pi^*(dx, i) \propto \frac{h_{1/i}(x)}{\theta^*(i)} 1_{X_1}(x)\lambda_1(dx)$. 

This can be done with Algorithm 2.2.

**Example 1.** We compare a plain simulated tempering with weight $\theta(i) \equiv 1$ and the Wang-Landau simulated tempering described above for sampling from

a multimodal bivariate Gaussian mixture distribution. The target distribution
given below was taken from Liang and Wong (2001)

\[
\pi(x) = \frac{1}{2\pi\sigma^2} \sum_{i=1}^{20} \omega_i \exp \left\{ - \frac{1}{2\sigma^2}(x - \mu_i)'(x - \mu_i) \right\},
\]

(3.4)

where $\sigma = 0.1$ and $\omega_i \equiv 0.05$. The $\mu_i$’s are listed in Table 1. The distribution is

highly multimodal and it is clear that a plain Random Walk Metropolis algorithm for

this distribution does not mix in a reasonable time. Simulated tempering can be particularly efficient in such situations. We compare two strategies: a plain

simulated tempering where $\theta(i) \equiv 1$ in (3.2) and the WL adaptation of simulated
tempering as described above. We use the temperature scale $1 < 7.7 < 31.6 < 100$.

Table 2 presents the mean squared errors (MSEs) of the two methods in

estimating the first two moments of the two components of $\pi$. We can see that

the WL version is about three-four times more efficient than the plain version

in terms of MSE. The estimates are based on 30 independent replications of the

samplers. We ran each sampler for 100,000 iterations. In applying Algorithm

2.2, we used $\gamma_n = 1/n$ and $c = 0.3$. 


Table 1. The 20 means of the two-dimensional Gaussian mixture.

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<th>i</th>
<th>µ₁₁</th>
<th>µ₁₂</th>
<th>i</th>
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Table 2. Mean Square Errors of the plain and the WL simulated tempering algorithms. Based on 30 independent replications with 100,000 iterations of each sampler.

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<td>WL-ST</td>
<td>0.029</td>
<td>0.041</td>
<td>2.818</td>
<td>4.023</td>
</tr>
<tr>
<td>Ratio</td>
<td>3.89</td>
<td>3.25</td>
<td>3.97</td>
<td>3.11</td>
</tr>
</tbody>
</table>

3.3. Application to trans-dimensional MCMC

It is often the case in Statistics that many alternative models are considered for the same data. One is then interested in issues like model comparison, model selection, and averaging. Let \( f(Data|k, x_k) \) be the likelihood of model \( k \) with parameter \( x_k \). Assume that we have a finite number \( d \) of models and that \( x_k \in (\mathcal{X}_k, \mathcal{B}_k, \lambda_k) \). Let \( \mathcal{X} = \bigcup_{i=1}^{d} \mathcal{X}_i \times \{ i \} \) be the union space equipped as above with the \( \sigma \)-algebra \( \mathcal{B} \) and the \( \sigma \)-finite measure \( \lambda \). \((\mathcal{X}, \mathcal{B}, \lambda)\) is the natural space to consider when dealing both with model uncertainty and parameter estimation.

In the Bayesian framework, a prior density (with respect to \( \lambda \)) \( p(x_k, k) \) in \((\mathcal{X}, \mathcal{B})\) is specified for \((x_k, k)\). The posterior distribution of \((x_k, k)\) is therefore \( \pi(x_k, k) \propto h_k(x_k) = f(Data|k, x_k)p(x_k, k) \). In this framework, one is often interested in the Bayes factor of model \( i \) to model \( j \) defined as \( B_{ij} := \theta^*(i)p(j)/\theta^*(j)p(i) \), where \( \theta^*(i) = \int_{\mathcal{X}_i} \pi(x, i)\lambda_i(dx_i) \) and \( p(i) = \int p(x, i)\lambda_i(dx_i) \). Trans-dimensional MCMC is a set of specialized MCMC algorithms to sample from distributions like \( \pi \) defined on spaces of variable dimensions. The reversible-jump algorithm of Green (Green 1995) is the most popular such sampler.

In the spirit of the WL algorithm, an alternative to sampling directly from \( \pi \) is to sample from the distribution

\[
\pi^*(dx_i, i) \propto \frac{h_i(x_i)}{\theta^*(i)}1_{\mathcal{X}_i}(x_i)\lambda_i(dx_i).
\]

By such re-weighting, we give the same posterior weight to all the models. The WL algorithm then offers an effective strategy to sample from \( \pi^* \) and we recover \( \pi \) by importance sampling. This strategy can improve on the mixing of the sampler.
Example 2. We set $\mathcal{X}_i = \mathbb{R}^i$ for $i = 1, \ldots, 20$, and consider the following rather trivial trans-dimensional target distribution:

$$
\pi(x_i, i) \propto a_i^{-1} e^{-|x_i|^2/2},
$$

where we let $a_i = 1$ for $i \neq 4$, and $a_4 = (2\pi)^{-16/2}$. In this distribution, $x_i \in \mathbb{R}^i$, the $i$-dimensional Euclidean space, and $\pi(x_i, i)$ restricted to $\mathbb{R}^i$ is proportional to the standard normal distribution. We are interested in the marginal distribution $p(i)$ of $i$. This distribution, as shown in Figure 1, is bimodal with modes at 4 and 20.

We pretend that this distribution is intractable and sample from it using a Birth-and-Death Reversible-Jump MCMC. For the fixed-dimensional move, we use a Random Walk Metropolis kernel with a Gaussian proposal with covariance matrix $\sigma_p I$, $\sigma_p = 0.1$. We implement a Birth-and-Death move for the trans-dimensional jump. Given $(x, i)$, we randomly select $j \in \{i - 1, i + 1\}$ with
respective probability $\omega_{i,i-1}, \omega_{i,i+1}$. We choose $\omega_{i,i+1} = 1/2$ with the usual correction at the boundaries. If $j = i+1$, we propose $y = (x_i, u)$ where $u \sim N(0, \sigma^2)$ with $\sigma = 0.1$. We accept $(y, j)$ with probability $\min(1, A)$, where

$$A = \frac{\pi(y, j) \omega_{ji}}{\pi(x, i) \omega_{ij}} \sqrt{2\pi\sigma} e^{u^2/2\sigma^2}.$$  

Similarly if $j = i-1$, we write $x = (y, u')$ with $u' \in \mathbb{R}$ and propose $(y, j)$. This value is then accepted with probability $\min(1, A)$ with

$$A = \frac{\pi(y, j) \omega_{ji}}{\pi(x, i) \omega_{ij}} \frac{1}{\sqrt{2\pi\sigma}} e^{-u'^2/2\sigma^2}.$$  

This vanilla RJMCMC sampler fails to sample from $\pi$. Depending on its starting point, the sampler typically found one of the two modes and got stuck there even after 10 millions iterations (see Figure 2a). In contrast, the WL algorithm provided a reasonable estimate of the distribution in only 2 millions iterations. In this example, computationally, each WL iteration step costs roughly that for 1.2 step of the vanilla sampler. For the WL approach we used $c = 0.4$ and $\gamma_n = 2^{-n}$ until $10^{-4}$ before switching to $\gamma_n = n^{-1}$.

4. Some Theoretical Results

We look at some theoretical aspects of the algorithm. We investigate the convergence of $\theta_n$ to $\theta^*$ and a strong law of large numbers for $\{(X_n, I_n)\}$. The difficulty comes in proving that the algorithm is stable in the following sense.

Definition 4.1. Let $v_n(i)$ be the occupation measure of $X_i$ by time $n$: $v_n(i) = (1/n) \sum_{k=1}^n 1_{I_k=i}$. The Wang-Landau algorithm is said to be stable if

$$\max_{i,j} \limsup_{n \to \infty} n(v_n(i) - v_n(j)) < \infty, \text{ a.s..} \quad (4.1)$$

This is an essential property of the algorithm. When the algorithm is stable, we show that all the stopping times $\kappa_l$ defined in Algorithm 2.2 are finite and the step-size $\gamma_n$ will gradually converges to 0. Moreover (Theorem 4.1 below) on a path where the algorithm is stable, $\theta_n \to \theta^*$ and a strong law of large numbers holds for $(1/n) \sum_{k=1}^n f(X_k, I_k)$. Then, we derive some verifiable conditions under which the algorithm is shown to be stable. To maintain the flow of ideas, some of the proofs are postponed to Section 6.

4.1. Ergodicity

Let $\Theta = \{ \theta \in \mathbb{R}^d : \sum_{i=1}^d \theta(i) = 1, \ \theta(i) \in (0, 1), \ i = 1, \ldots, d \}$ and $((X_0, I_0), \theta_0) \in \mathcal{X} \times \Theta$ be the initial state of the algorithm. This initial state
is considered fixed but arbitrary. Let \( \{ \gamma_n \} \) be the step-size sequence. Let \( \text{Pr} \) be the distribution of the process \( \{ X_n, I_n, \theta_n \} \) started at \( (X_0, I_0, \theta_0) \) with step-size sequence \( \{ \gamma_n \} \), and let \( \Omega \) be the expectation with respect to \( \text{Pr} \). To simplify the notation, we omit explicit mention of the dependence of \( \text{Pr} \) on \( (X_0, I_0, \theta_0) \) and \( \{ \gamma_n \} \). All statements made almost surely are with respect to \( \text{Pr} \). For \( \theta \in \mathbb{R}^d \), \( |\theta| \) denotes the Euclidean norm of \( \theta \). For any \( \varepsilon \in (0, \varepsilon_*) \), where \( \varepsilon_* = \min_i \theta^*(i) \), let \( \Theta_{\varepsilon} = \{ \theta \in \Theta, \, \theta(i) \geq \varepsilon, \, i = 1, \ldots, d \} \). Our main assumption is that the family \( \{ P_\theta, \theta \in \Theta_{\varepsilon} \} \) is Lipschitz and uniformly \( V \)-ergodic. See e.g., [Andrieu and Moulines (2006)] for some examples of MCMC samplers where these assumptions hold. Before stating them, we need some notation. For any functions \( f : \mathcal{X} \to \mathbb{R} \) and \( W : \mathcal{X} \to [1, \infty) \) we take \( |f|_W := \sup_{x \in \mathcal{X}}(|f(x)|/W(x)) \), and introduce the set of \( W \)-bounded functions \( \mathcal{W} := \{ f \text{ meas.}, \, f : \mathcal{X} \to \mathbb{R}, \, |f|_W < \infty \} \). A transition kernel on \( (\mathcal{X}, \mathcal{B}) \) operates on measurable real-valued functions \( f \) as \( Pf(x) = \int P(x, dy)f(y) \), and the product of two transition kernels \( P_1 \) and \( P_2 \) is the transition kernel \( P_1P_2(x, A) = \int P_1(x, dy)P_2(y, A) \). For two transition kernels \( P_1 \) and \( P_2 \), \( \|P_1 - P_2\|_W \), the \( W \)-distance between \( P_1 \) and \( P_2 \) is
\[
\|P_1 - P_2\|_W := \sup_{|f|_W \leq 1} |P_1f - P_2f|_W.
\]

(A1) There is a measurable function \( V : \mathcal{X} \to [1, \infty) \), a set \( C \subset \mathcal{X} \), and a probability measure \( \nu \) on \((\mathcal{X}, \mathcal{B})\) such that \( \nu(C) > 0 \), with the following property. For all \( \varepsilon \in (0, \varepsilon_*) \), we can find constants \( \lambda_\varepsilon \in (0, 1) \), \( b_\varepsilon \in [0, \infty) \), \( \beta_\varepsilon \in (0, 1] \) and integer \( n_{0, \varepsilon} \) such that:
\[
\inf_{\theta \in \Theta_{\varepsilon}} P_{\theta}^{n_{0, \varepsilon}}(x, A) \geq \beta_\varepsilon \nu(A)1_C(x), \quad x \in \mathcal{X}, \quad A \in \mathcal{B};
\]
\[
\sup_{\theta \in \Theta_{\varepsilon}} P_\theta V(x) \leq \lambda_\varepsilon V(x) + b_\varepsilon 1_C(x), \quad x \in \mathcal{X}.
\]

The inequality (4.3) of (A1) is the so-called drift condition and (4.2) is the so-called minorization condition.

(A2) For all \( \alpha \in [0, 1] \), for all \( \varepsilon \in (0, \varepsilon_*) \), there exists \( K = K(\alpha, \varepsilon) < \infty \) such that, for all \( \theta, \, \theta' \in \Theta_{\varepsilon} \),
\[
\|P_{\theta} - P_{\theta'}\|_{V_\alpha} \leq K |\theta - \theta'|,
\]
where \( V \) is defined in (A1).

(A3) \( \{ \gamma_n \} \) is non-increasing, \( \gamma_n > 0 \), \( \sum \gamma_n = \infty \), and \( \gamma_n = O(n^{-1}) \) as \( n \to \infty \).
For $B > 0$, we introduce the stopping time:
\[
\tau(B) := \inf \left\{ k \geq 0 : \max_{i,j} k (v_k(i) - v_k(j)) > B \right\},
\]
with the usual convention that $\inf \emptyset = \infty$. $\tau(B)$ is the first time that one component accumulates $B$ visits or more than some other component. Thus, Definition 4.1 is precisely equivalent to $\tau(B) = \infty$ for some $B$. With this in mind, the next results says essentially that if the algorithm is stable, and (A1-3) hold, then it is ergodic.

**Theorem 4.1.** Let (A1–3) hold and $B > 0$ be given. Then

(i) \[|\theta_n - \theta^*| 1_{\{\tau(B) > n\}} \to 0, \text{ a.s. as } n \to \infty.\] (4.6)

(ii) For any function $f \in L_{V^{1/2}}$, writing $\bar{f} = f - \pi^*(f)$, we have
\[
\frac{1}{n} n \sum_{k=1}^n \bar{f}(X_k, I_k) 1_{\{\tau(B) > k-1\}} \to 0, \text{ a.s. as } n \to \infty.
\] (4.7)

**Proof.** See Section 6.2.

4.2. Checking (A1–2)

We impose the drift condition and the minorization condition (A1) uniformly for $\theta \in \Theta_\varepsilon$, not uniformly for $\theta \in \Theta$. This is an important point. Indeed, and as we see now, a drift and minorization condition uniformly-in-$\theta$ for $\theta \in \Theta_\varepsilon$ is almost always true as soon as one $P_\theta$ satisfies these conditions (Proposition 4.1), whereas a minorization and drift condition uniformly-in-$\theta$ for $\theta \in \Theta$ is almost never true.

Indeed, suppose that each $P_\theta$ is a Metropolis-Hastings kernel with invariant distribution $\pi_\theta$ and proposal kernel $Q$. That is $P_\theta f(x, i) = M_\theta f(x, i) + r_\theta(x, i) f(x, i)$, where

\[
M_\theta f(x, i) = \sum_{j=1}^d \int_{X_j} \min \left( 1, \frac{\theta(i)}{\theta(j)} R(y, j; x, i) \right) f(y, j) Q(x, i; dy, j),
\]

\[
r_\theta(x, i) = 1 - \sum_{j=1}^d \int_{X_j} \min \left( 1, \frac{\theta(i)}{\theta(j)} R(y, j; x, i) \right) Q(x, i; dy, j),
\]

and $R(x, i; y, j)$ the Radon-Nikodym density of $\pi(dx, i)Q(x, i; dy, j)$ with respect to $\pi(dy, j)Q(y, j; dx, i)$. 

With \( \theta = (1, \ldots, 1) \), we use \( \pi \) (resp. \( P \) and \( M \)) to denote \( \pi_\theta \) (resp. \( P_\theta \) and \( M_\theta \)). The next result states that we only need to check that \( P \) is geometrically ergodic to obtain (A1-2).

**Proposition 4.1.** Suppose there is a measurable function \( V : \mathcal{X} \rightarrow [1, \infty) \), a set \( C \subset \mathcal{X} \), a probability measure \( \nu \) on \((\mathcal{X}, \mathcal{B})\) such that \( \nu(C) > 0 \), and constants \( \lambda \in (0, 1) \), \( b \in [0, \infty) \), \( \beta \in (0, 1] \) and finite integer \( n_0 \) such that \( \nu(x) \geq \beta \nu(A) \mathbf{1}_C(x) \), \( x \in \mathcal{X} \), \( A \in \mathcal{B} \), and \( PV(x) \leq \lambda V(x) + b \mathbf{1}_C(x) \), \( x \in \mathcal{X} \). Then (A1-2) hold.

**Proof.** See Section 6.1.

### 4.3. Stability

Theorem 4.1 asserts that under (A1–3), the WL algorithm converges to the right limit on stable paths. This raises the question of checking the stability of the algorithm, difficult in general. The next theorem gives some easily checked conditions under which the algorithm is stable.

**Theorem 4.2.** The WL algorithm is stable under either of the following two conditions.

(a) There exist \( \varepsilon \in (0, 1) \), \( K \in (0, \infty) \) and integer \( n_0 \geq 0 \) such that for any \( i, j \in \{1, \ldots, d\} \) and \( \theta \in \mathbb{R}^d \), \( \theta(i)/\theta(j) > K \) implies that \( P_\theta^{n_0}((x, i), \mathcal{X}_j \times \{j\}) \geq \varepsilon \) for all \( x \in \mathcal{X}_i \).

(b) There exists \( \varepsilon \in (0, 1) \), such that for any \( j \in \{1, \ldots, d\} \) and \( \theta \in \mathbb{R}^d \), \( \theta(j) \leq \min_{1 \leq i \leq d} \theta(i) \) implies \( P_\theta((x, i), \mathcal{X}_j \times \{j\}) \geq \varepsilon \) for all \( x \notin \mathcal{X}_i \) and all \( i \neq j \).

**Proof.** See Section 6.3.

### 4.4. Application to multicanonical sampling

Consider the multicanonical sampling of Section 3.1. Suppose that \( P_\theta \) is the Independence-Sampler with proposal distribution \( Q(dx) = q(x)\lambda(dx) \) and invariant distribution \( \pi_\theta(dx) \propto \sum_{i=1}^d (h(x)/\theta(i)) \mathbf{1}_{\mathcal{X}_i}(x) \lambda(dx) \). Assume that the function \( \omega(x) \propto h(x)/q(x) \) is bounded with supremum \( \omega_0 \). Then clearly, for all \( x \in \mathcal{X}_i \),

\[
P_\theta(x, \mathcal{X}_j) \geq \min \left( 1, \frac{\theta(i)}{\theta(j)} \right) \int_{\mathcal{X}_j} \min \left( 1, \frac{\omega(y)}{\omega_0} \right) Q(dy) \geq \varepsilon_j,
\]

as soon as \( \theta(i) \geq \theta(j) \), taking \( \varepsilon_j = \int_{\mathcal{X}_j} \min(1, [(\omega(y))/(\omega_0)]) Q(dy) > 0 \) (since \( \pi(\mathcal{X}_i) > 0 \)). Thus for any \( i, j \in \{1, \ldots, d\} \), \( i \sim j \) and, by Theorem 4.2, the WL algorithm is stable. Now, since \( \omega \) is bounded, each \( P_\theta \) satisfies a drift condition and a minorization condition, which implies (A1–2) by Proposition 4.1.
Similarly, if $P_\theta$ is a Random Walk Metropolis with proposal kernel $q(y-x)$ we have

$$P_\theta(x, x_j) \geq \min \left( 1, \frac{\theta(i)}{\theta(j)} \right) \int_{x_j} \min \left( 1, \frac{\pi(y)}{\pi(x)} \right) q(y-x) dy.$$ 

It follows that if $X$ is compact and $\pi$, $q$ are positive and continuous, then $i \sim j$ for all $i, j$. Under the same assumption, (A1–2) also hold.

**Corollary 4.1.** In the case of multicanonical sampling of Section 3.1, assume one of

(i) $P_\theta$ is an independent-Metropolis sampler with proposal distribution $Q(dx) = q(x)\lambda(dx)$ and $\omega \propto h/q$ is bounded;

(ii) $P_\theta$ is a RWM sampler with proposal $q(y-x)$; $X$ is compact, and $\pi$ and $q$ are positive and continuous.

Then the algorithm is stable and, under (A3),

$$|\theta_n - \theta^*| \to 0; \quad \text{and} \quad \frac{1}{n} \sum_{k=1}^n f(X_k) \to \pi^*(f) \quad \text{a.s. as} \quad n \to \infty$$

for any bounded measurable function $f$.

**4.5. Application to simulated tempering**

Theorems 4.1. and 4.2. can also be applied to the WL version of simulated tempering as described in Section 3.2. We consider the simulated tempering algorithm with kernel $P_\theta((x, i); A \times \{j\}) = B_{x, \theta}(i, j)P^{[j]}(x, A)$ where $B_{x, \theta}(i, j) \propto h_j(x)/\theta(j)$ and $P^{[j]}$ is a transition kernel (not necessarily Metropolis-Hastings) with invariant distribution $h_j$. The following corollary is easily proved and omitted.

**Corollary 4.2.** Suppose that $X_1$ is compact, $h$ positive and continuous, and $\{\gamma_n\}$ satisfies (A3). Suppose that there exist $\varepsilon > 0$, a probability measure $\nu$ and an integer $n_0$ such that $\nu(X_i) \geq 0$ and $(P^{[i]})^{n_0}(x, A) \geq \varepsilon \nu(A)$, $i \in \{1, \ldots, d\}$. Then

$$|\theta_n - \theta^*| \to 0; \quad \text{and} \quad \frac{1}{n} \sum_{k=1}^n f(X_k, I_k) \to \pi^*(f) \quad \text{a.s. as} \quad n \to \infty$$

for any bounded measurable function $f$.

**5. Discussion and Open Problems**

In this paper, we propose an extension of the WL algorithm to general state spaces. The WL algorithm differs from other adaptive Markov Chain Monte
Carlo algorithms based on stochastic approximation by the adaptive nature of its step-size. We have shown through examples that the algorithm can be used effectively to improve on simulated tempering and trans-dimensional MCMC algorithms. We have also studied the asymptotic behavior of the WL algorithm. We have shown that on stable sample paths, and with an appropriate step-size, \( \theta_n \) converges to \( \theta^* \) and a strong law of large numbers holds. Finally, when the state space is compact we have shown that, in most cases, the algorithm is stable and the aforementioned limit results apply.

Two main questions have remained unanswered. The first concerns the stability of the algorithm in unbounded state spaces. Second, in order to exploit the full potential of the algorithm, a more precise understanding of its efficiency is needed. In particular, we need to understand how the rate of convergence and the asymptotic variances of the algorithm are related to the parameter \( c \) and step-size \( \gamma_n \). We are currently investigating some of these questions.

6. Proofs

The techniques used here can be found in various forms in [Benveniste, Métivier and Priouret (1990), Delyon (1996), and Andrieu, Moulines and Priouret (2005)]. Throughout, \( C_\epsilon \) denotes a generic constant whose value can be different from one equation to another. The key result in the proof of Theorem 4.1 is Lemma 6.6, which states that the weighted sum of the noise process in the stochastic approximation followed by \( f \theta_n \) is summable.

6.1. Proof of Proposition 4.1

**Lemma 6.1.** Under the conditions of Proposition 4.1, (A2) holds.

**Proof.** Let \( \varepsilon \in (0, \varepsilon^*) \) and \( \alpha \in (0, 1] \). For \( \theta \in \Theta_\varepsilon \), \( |f| \leq V^\alpha \), and \( (x, i) \in \mathcal{X} \), we have \( P_\theta f(x, i) = M_\theta f(x, i) + f(x, i)(1 - M_\theta e(x, i)) \), where \( e(x, i) \equiv 1 \), and

\[
M_\theta f(x, i) = \sum_{j=1}^{d} \min_{\mathcal{X}} \left( 1, \frac{\theta(i)}{\theta(j)} r(y, j; x, i) \right) f(y, j) Q(x, i; dy, j).
\]

As a consequence, for \( \theta_2, \theta_1 \in \Theta_\varepsilon \), \( \|P_{\theta_2} - P_{\theta_1}\|_{V^\alpha} \leq \|M_{\theta_2} - M_{\theta_1}\|_{V^\alpha} + \|M_{\theta_2} - M_{\theta_1}\|_{TV} \), where \( \|P\|_{TV} = \|P\|_V \) with \( V \equiv 1 \). For \( (x, i) \in \mathcal{X} \), the function \( \theta \rightarrow M_{\theta} f(x, i) \) is differentiable and:

\[
\sum_{j=1}^{d} \left| \frac{\partial}{\partial \theta_j} M_{\theta} f(x, i) \right| \leq C_\epsilon \sum_{j=1}^{d} M |f| (x, i).
\]

(A2) then follows by the Mean Value Theorem and the drift condition on \( P \).
Next we show that for any $\varepsilon \in (0, \varepsilon^*)$, the family $(P_\theta)_{\theta \in \Theta_\varepsilon}$ satisfies a uniform (in $\theta$) drift condition.

**Lemma 6.2.** Under the conditions of Proposition 4.1, (A1) holds.

**Proof.** The minorization is immediate. Since $\min(1, ab) \geq \min(1, a) \min(1, b)$ for $a, b > 0$, $P_\theta(x, i; A) \geq M_\theta(x, i; A) \geq (1/\varepsilon) M_\theta(x, i; A) \geq \beta/\varepsilon \nu(A) 1_C(x, i)$.

The argument to show the uniform drift is fairly simple, and we only sketch it here. Let $B_{\theta, r} = \Theta_\varepsilon \cap B(\theta, r)$ the open ball of $\Theta_\varepsilon$ with center $\theta \in \Theta_\varepsilon$ and radius $r > 0$. It follows from Lemma 6.1 that by choosing $r > 0$ small enough, if $P_\theta$ satisfies a drift condition toward a small set $C$, then the family $(P_{\theta'}, \theta' \in B_{\theta, r})$ satisfies a uniform (in $\theta'$) drift condition toward $C$. Therefore, starting from $P$, we can find an open coverage of $\Theta_\varepsilon$ by the $B_{\theta, r}$ such that, on each such ball, a uniform drift toward $C$ holds. Since $\Theta_\varepsilon$ is compact it admits a finite coverage by the $B_{\theta, r}$ and, using the maximum of the constants of the drift condition toward $C$ for each such ball, we get a uniform drift toward $C$ for $(P_\theta, \theta \in \Theta_\varepsilon)$.

### 6.2. Proof of Theorem 4.1.

We start with some preliminary remarks. For any $\varepsilon \in (0, \varepsilon^*)$ and $\alpha \in (0, 1]$, it is known from Markov chain theory that (A1) implies the existence of $C_{\varepsilon, \alpha} < \infty$, $\rho_{\varepsilon, \alpha} \in (0, 1)$, and $b_\varepsilon$ as in (A1), such that

$$\sup_{\theta \in \Theta_\varepsilon} \| P^n_{\theta} - \pi_\theta \|_{V^\alpha} \leq C_{\varepsilon, \alpha} \rho_{\varepsilon, \alpha}^n, \quad (6.1)$$

$$\sup_{\theta \in \Theta_\varepsilon} \pi_\theta(V) \leq b_\varepsilon. \quad (6.2)$$

For a proof, see e.g., Baxendale (2005) and the references therein. Define $\xi(\varepsilon) = \inf\{k \geq 0 : \theta_k \notin \Theta_\varepsilon\}$. An easy calculation using (A1) gives that

$$E \left[ V(X_n, I_n) 1_{\{\xi(\varepsilon) > n\}} \right] = E \left[ V(X_n, I_n) 1_{\Theta_\varepsilon}(\theta_0) \cdots 1_{\Theta_\varepsilon}(\theta_n) \right]$$

$$\leq \lambda^n V(X_0, I_0) + \frac{b_\varepsilon}{(1 - \lambda)}$$

from which we deduce that

$$\sup_n E \left[ V(X_n, I_n) 1_{\{\xi(\varepsilon) > n\}} \right] < \infty. \quad (6.3)$$

The proof of the theorem is based on some nice properties of solutions of the so-called Poisson equation. These solutions allow us to obtain a martingale approximation to the process $\sum_{k=1}^n f(X_k, I_k)$. For $f \in L_{V^\alpha}$, $\alpha \in (0, 1]$ and $\theta \in \Theta_\varepsilon$, let

$$h_\theta = \sum_{k=0}^\infty P^k_{\theta} \left( f - \pi_\theta(f) \right). \quad (6.4)$$
that using (A1), \( h_\theta \) exists and \( h_\theta \in L_{V^\alpha} \). For \( f \in L_{V^\alpha} \) and \( \theta, \theta' \in \Theta_\varepsilon \), writing \( \bar{f}_\theta = f - \pi_\theta(f) \), we have

\[
|\pi_\theta(f) - \pi_{\theta'}(f)| = |\pi_\theta\left[P^k_\theta \bar{f}_{\theta'}\right]|
\]

\[
= |\pi_\theta\left[P^k_\theta \bar{f}_{\theta'}\right] + \sum_{j=1}^k \pi_\theta\left[P^{k-j}_\theta (P_\theta - P_{\theta'})P^{j-1}_{\theta'}(\bar{f}_{\theta'})\right]|
\]

\[
\leq C_\varepsilon \left(\rho_\varepsilon^2 + |\theta - \theta'| \sum_{j=1}^k \rho_\varepsilon^{j-1}\right),
\]

using (A1–2), from which we deduce that there exists a finite constant \( C_\varepsilon \) such that

\[
\sup_{f \in L_{V^\alpha}} |\pi_\theta(f) - \pi_{\theta'}(f)| \leq C_\varepsilon |\theta - \theta'|. \tag{6.5}
\]

The constant \( C_\varepsilon \) is not necessarily the same from one equation to the other. Similarly, with \( \bar{P}_\theta \) as the operator \( P_\theta - \pi_\theta \), we have

\[
|P^k_\theta \bar{f}_\theta - P^k_{\theta'} \bar{f}_{\theta'}| = |P^k_\theta f - P^k_{\theta'} f|
\]

\[
= \left|\sum_{j=1}^k \bar{P}^{k-j}_\theta (P_\theta - P_{\theta'})\bar{P}^{j-1}_{\theta'} f\right|
\]

\[
\leq C_\varepsilon |\theta - \theta'| k\rho_\varepsilon^{k-1}V^\alpha,
\]

using (A1–2). This, together with (6.3), implies the existence of a finite constant \( C_\varepsilon \) such that for all \( \alpha \in (0, 1] \), \( \theta, \theta' \in \Theta_\varepsilon \),

\[
|h_\theta - h_{\theta'}|_{V^\alpha} + |P_\theta h_\theta - P_{\theta'} h_{\theta'}|_{V^\alpha} \leq C_\varepsilon |\theta - \theta'|. \tag{6.6}
\]

In our analysis, we mainly see \( \{\theta_n\} \) as a stochastic approximation sequence. The recursion on \( \{\theta_n\} \) has

\[
\theta_{n+1}(i) = \frac{\phi_{n+1}(i)}{\sum_{e=1}^d \phi_{n+1}(e)}
\]

\[
= \frac{\phi_n(i) + \gamma_{\alpha_n} \phi_n(i)1_{\{I_{n+1} = i\}}}{\sum_{e=1}^d \phi_n(e) + \gamma_{\alpha_n} \phi_n(I_{n+1})}
\]

\[
= \frac{\theta_n(i) \left[1 + \gamma_{\alpha_n} \frac{1}{\sum_{e=1}^d \phi_n(e) + \gamma_{\alpha_n} \phi_n(I_{n+1})}\right]}{1 + \gamma_{\alpha_n} \theta_n(I_{n+1})}
\]

\[
= \theta_n(i) + \gamma_{\alpha_n} H_{i}(\theta_n, I_{n+1}) + \gamma_{\alpha_n}^2 r_{i,n}(\theta_n, I_{n+1}), \tag{6.7}
\]

where \( H_{i}(\theta, I) = \theta(i)\left[1_{\{I = i\}} - \theta(I)\right] \) and \( r_{i,n}(\theta, I) = -\theta(i)\theta(I)\right[\left[1_{\{I = i\}} - \theta(I)\right]/\left(1 + \gamma_{\alpha_n} \theta(I)\right)\right]\).
The mean field function $h_i(\theta) = \pi_\theta(H_i)$ is
\[
h_i(\theta) = \frac{\theta^*(i) - \theta(i)}{\sum_{j=1}^{d} \theta^*(j)}/\theta(j). \quad (6.8)
\]

**Lemma 6.3.** Under (A3), let $B > 0$ be given and $c_0 = 2Bd/c$. Then on \{\(\tau(B) > n\)\}, $a_n \geq [n/c_0]$. Moreover $\sum \gamma_{a_n} = \infty$, and $\sum \gamma_{a_n}^2 1_{\{\tau(B) > n\}} < \infty$ almost surely.

**Proof.** On \{\(\tau(B) > n\)\}, for any $k < k' \leq n$, and for any $i \in \{1, \ldots, d\}$, $(v_{k,k'}(i) - 1/d) \leq 2B/(k' - k)$. Therefore if $l_{c_0} \leq n < \tau(B)$, $\kappa_l \leq n$. That is, $a_n \geq [n/c_0]$ on \{\(\tau(B) > n\)\}.

Since \{\(\gamma_n\)\} is non-increasing and $a_n \leq n$, $\sum \gamma_{a_n} \geq \sum \gamma_n = \infty$. On the other hand $\sum \gamma_{a_n}^2 1_{\{\tau(B) > n\}} \leq C0 \sum n^{-2} < \infty$.

For any $\varepsilon > 0$, we introduce the stopping time $\xi(\varepsilon) := \inf\{k \geq 0 : \theta_k \not\in \Theta_{\varepsilon}\}$. We need the following result. The proof is omitted.

**Lemma 6.4.** Let \{\(\gamma_n, n \geq 0\)\} be a non-increasing sequence of positive number and \{\(v_n, n \geq 0\)\} be a sequence of numbers such that $|\sum_{k=0}^{N} v_k| \leq B$ for all $N \geq 0$. Then $|\sum_{k=0}^{N} \gamma_k v_k| \leq \gamma_0 B$ for all $N \geq 0$.

The following lemma relates $\tau(B)$ and $\xi(\varepsilon)$.

**Lemma 6.5.** Under (A3), for any $B > 0$ we can find $\varepsilon \in (0, \varepsilon^*)$ such that $\tau(B) \leq \xi(\varepsilon)$.

**Proof.** Take $\varepsilon = (1 + (d - 1)e^{B\theta})^{-1} > 0$. Without any loss, we assume that $\varepsilon < \varepsilon^*$ and $\varepsilon \leq \min\theta(1)$. We need to show that $\min\theta(i) > \varepsilon$ for all $\tau_0 < \tau(B)$. But $\theta(1) = (1 + \sum_{j \neq i}(\phi(i))/\phi(i)))^{-1}$. It is thus enough to show that $\phi(i)/\phi(i) \leq e^{B\theta}$ for all $i \neq j$ and for any $n < \tau(B)$. But
\[
\frac{\phi(i)}{\phi(i)} = \exp\left(\sum_{i=0}^{N} \gamma_p(N_{m,n,n+1}(i) - N_{m,n,n+1}(i))\right),
\]
where $N_{m,n}(i) = 0$ if $m \leq 1$ and $N_{m,n}(i) = \sum_{m=1}^{N} \mathcal{X}_t(X_n)$ otherwise ($N_{m,n}(i)$ is the number of visits to $X$ from time $1$ to $m$). For any $n < \tau(B)$, $|\sum_{p=0}^{N} (N_{m,n,n+1}(i) - N_{m,n,n+1}(i))| \leq B$ for all $P \geq 0$. Lemma 6.4 thus implies that $\phi(i)/(\phi(j) \leq e^{B\theta}$.

**Lemma 6.6.** Under (A1–3), let $B > 0$ be given. Let \{\(\gamma_n\)\} be a sequence that satisfies (A3) such that $\sum \gamma_{[n/a]} \gamma_{[n/a]} < \infty$ for all $a > 0$. For $\theta \in \Theta$, Let $H_\theta : \mathcal{X} \to \mathbb{R}$ be a measurable function such that $H_\theta \in L_{V/2}$. Then
\[
\sum_{k=0}^{\infty} \gamma_{a_k}^2 1_{\{\tau(B) > k\}} \left[H_{\theta_k}(X_{k+1}, 1_{k+1}) - \pi_{\theta_k}(H_{\theta_k})\right] < \infty, \text{ a.s.} \quad (6.9)
\]
Proof. Let $\varepsilon > 0$ as in Lemma 6.5. From (A1) there is a function $h_\theta \in L_{V^{1/2}}$ that solves the Poisson equation $h_\theta - P_0 h_\theta = H_\theta - \pi_\theta (H_\theta)$ for all $\theta \in \Theta_\varepsilon$. Using this we can write

$$\sum_{k=0}^{n} \gamma'_a 1_{\{\tau(B)>k\}} \left[ H_{\theta_k}(X_{k+1}, I_{k+1}) - \pi_{\theta_k}(H_{\theta_k}) \right]$$

$$= \sum_{k=0}^{n} \gamma'_a 1_{\{\tau(B)>k\}} \left( U^{(1)}_{k+1} + U^{(2)}_{k+1} + U^{(3)}_{k+1} \right),$$

where

$$U^{(1)}_{k+1} = h_{\theta_k}(X_{k+1}, I_{k+1}) - P_{\theta_k} h_{\theta_k}(X_k, I_k),$$

$$U^{(2)}_{k+1} = P_{\theta_k} h_{\theta_k}(X_k, I_k) - P_{\theta_{k+1}} h_{\theta_{k+1}}(X_{k+1}, I_{k+1}),$$

$$U^{(3)}_{k+1} = P_{\theta_{k+1}} h_{\theta_{k+1}}(X_{k+1}, I_{k+1}) - P_{\theta_k} h_{\theta_k}(X_k, I_k).$$

Clearly $M_n = \sum_{k=0}^{n} \gamma'_a 1_{\{\tau(B)>k\}} U^{(1)}_{k+1}$ is a martingale; using Lemma 6.3, (6.3) and since $h_\theta$ and $P_0 h_\theta \in L_{V^{1/2}}$, we have

$$E(M_n^2) \leq C \varepsilon \sum_{k=0}^{n} \left( \gamma'_a \right)^2 X_{k+1}, I_{k+1}$$

$$\leq C \varepsilon \left( \sum_{k=0}^{\infty} \left( \gamma'_a \right)^2 \right) E \left[ 1_{\{\tau(B)>k\}} V(X_{k+1}, I_{k+1}) \right]$$

$$\leq C \varepsilon \sum_{k=0}^{\infty} \left( \gamma'_a \right)^2 < \infty.$$

By Doob’s Convergence Theorem for martingales, $\sum_{k=0}^{\infty} \gamma'_a 1_{\{\tau(B)>k\}} U^{(1)}_{k+1}$ is finite a.s.

On $\{\tau(B) = l\}$ ($l < \infty$), $\sum_{k=0}^{\infty} \gamma'_a 1_{\{\tau(B)>k\}} U^{(2)}_{k+1} = \sum_{k=0}^{l-1} \gamma'_a U^{(2)}_{k+1}$, which is finite almost surely. On $\{\tau(B) = \infty\}$, we can write

$$1_{\{\tau(B) = \infty\}} \sum_{k=0}^{n} \gamma'_a 1_{\{\tau(B)>k\}} U^{(2)}_{k+1}$$

$$= \gamma'_a P_{\theta_0} h_{\theta_0}(X_0, I_0) - \gamma'_a 1_{\{\tau(B) = \infty\}} P_{\theta_{n+1}} h_{\theta_{n+1}}(X_{n+1}, I_{n+1})$$

$$+ \sum_{k=0}^{n-1} \left( \gamma'_a - \gamma'_a \right) 1_{\{\tau(B) = \infty\}} P_{\theta_{k+1}} h_{\theta_{k+1}}(X_{k+1}, I_{k+1}).$$
Now \( E[(\gamma_{a_k}^t 1_{\{\tau(B) = \infty\}} P_{\theta_{k+1}} h_{\theta_{k+1}}(X_{n+1}, I_{n+1}))^2] \leq C_\varepsilon (\gamma_{[n/c_0]}^t)^2 \) and \( \sum (\gamma_{[n/c_0]}^t)^2 < \infty \), thus \( \gamma_{a_k}^t 1_{\{\tau(B) = \infty\}} P_{\theta_{k+1}} h_{\theta_{k+1}}(X_{n+1}, I_{n+1}) \) converges a.s. to 0, and

\[
\sum_{k=0}^{n-1} |(\gamma_{a_{k+1}}^t - \gamma_{a_k}^t) 1_{\{\tau(B) = \infty\}} P_{\theta_{k+1}} h_{\theta_{k+1}}(X_{k+1}, I_{k+1})| \leq C_\varepsilon \sum_{k=0}^{\infty} (\gamma_{a_{k+1}}^t - \gamma_{a_k}^t) 1_{\{\tau(B) = \infty\}} V^{1/2}(X_{k+1}, I_{k+1}).
\]

Since \( a_k \) only changes at the stopping times \( \kappa_i \), we have

\[
E \left[ \sum_{k=0}^{\infty} (\gamma_{a_k}^t - \gamma_{a_{k+1}}^t) 1_{\{\tau(B) = \infty\}} V^{1/2}(X_{k+1}, I_{k+1}) \right] = E \left[ \sum_{k=1}^{\infty} (\gamma_{k-1}^t - \gamma_k^t) 1_{\{\tau(B) = \infty\}} V^{1/2}(X_{n_k+1}, I_{n_k+1}) \right]
\]

\[
= \sum_{k=1}^{\infty} (\gamma_{k-1}^t - \gamma_k^t) E \left[ 1_{\{\tau(B) = \infty\}} V^{1/2}(X_{n_k+1}, I_{n_k+1}) \right]
\]

\[
\leq C_\varepsilon \sum_{k=1}^{\infty} (\gamma_{k-1}^t - \gamma_k^t)
\]

\[
\leq C_\varepsilon \gamma_0.
\]

By Lebesgue’s Dominated Convergence Theorem, we can conclude that

\[
E \left[ \sum_{k=0}^{\infty} (\gamma_{a_{k+1}}^t - \gamma_{a_k}^t) 1_{\{\tau(B) = \infty\}} P_{\theta_{k+1}} h_{\theta_{k+1}}(X_{k+1}, I_{k+1}) \right] < \infty.
\]

This is sufficient to conclude that \( \sum_{k=0}^{\infty} \gamma_{a_k}^t 1_{\{\tau(B) > k\}} U^{(2)}_{k+1} \) is finite almost surely.

Using (6.5):

\[
\sum_{k=0}^{n} \gamma_{a_k}^t 1_{\{\tau(B) > k\}} U^{(3)}_{k+1} \leq C_\varepsilon \sum_{k=0}^{\infty} \gamma_{a_k}^t \gamma_{a_{k+1}} 1_{\{\tau(B) > k\}} V^{1/2}(X_{k+1}, I_{k+1})
\]

\[
\leq C_\varepsilon \sum_{k=0}^{\infty} \gamma_{[n/c_0]} \gamma_{[n/c_0]} 1_{\{\tau(B) > k\}} V^{1/2}(X_{k+1}, I_{k+1}),
\]

and

\[
E \left[ \sum_{k=0}^{\infty} \gamma_{[n/c_0]} \gamma_{[n/c_0]} 1_{\{\tau(B) > k\}} V^{1/2}(X_{k+1}, I_{k+1}) \right] \leq C_\varepsilon \sum_{k=0}^{\infty} \gamma_{[n/c_0]} \gamma_{[n/c_0]} < \infty
\]
by (6.3). With Lebesgue’s Dominated Convergence Theorem, we deduce that
\[ E \left[ \sum_{k=0}^{\infty} \gamma_{ak} 1_{\{\tau(B)>k\}} \left| U^{(3)}_{k+1} \right| \right] \leq C_{s} \sum_{k=0}^{\infty} \gamma_{[n/c]} \gamma_{[n/c]} < \infty, \]
which implies that \( \sum_{k=0}^{n} \gamma_{ak} 1_{\{\tau(B)>k\}} U^{(3)}_{k+1} \) converges almost surely to a finite limit. This completes the proof of the lemma.

We are now in position to prove Theorem 4.1. We start with (i).

**Proposition 6.1.** Under (A1–3), let \( B > 0 \) be given. Then \( \theta_n - \theta^* 1_{\{\tau(B)>n\}} \rightarrow 0 \) with probability one as \( n \rightarrow \infty \).

**Proof.** The idea of the proof is borrowed from [Delyon (1996)]. We saw in (6.7) that \( \theta_{n+1} = \theta_n + \gamma_{n} H(\theta_n, I_{n+1}) + \gamma_{2} n, \) where \( H = (H_1, \ldots, H_d), r_n = (r_{n,1}, \ldots, r_{n,d}), H_i(\theta, I) = \theta(i)(1_{i=i} - \theta(I)) \) and \( r_{i,n} = -\theta(i)\theta(I)(1_{i=i} - \theta(I))/(1 + \gamma_{n} \theta(I)). \) We note that \( |H| \leq 1 \) and \( |r_n| \leq 1. \) Let \( \varepsilon \in (0, \varepsilon^*) \) be such that \( \tau(B) \leq \xi(\varepsilon) \) (Lemma 6.5). We recall that the mean field function of the recursion is \( h_i(\theta) = (\theta^*(i) - \theta(i))/\sum_{j=1}^{d}(\theta^*(j))/(\theta(j)) \). We introduce \( \theta'_n = \theta_n + \gamma_{n} 1_{\{\tau(B)>j\}} [H(\theta_j, I_{j+1}) - h(\theta_j)]. \) From Lemma 6.6, \( |\theta'_n - \theta| \rightarrow 0 \) almost surely. \( \{\theta'_n\} \) satisfies the recursion
\[
\theta'_{n+1} = \theta'_n + \gamma_{n} h(\theta_n) + \gamma_{a} 1_{\{\tau(B)\leq n\}} \left[ H(\theta_n, I_{n+1}) - h(\theta_n) \right] + \gamma_{2} n, \]

We can then deduce that
\[
\left| \theta'_{n+1} - \theta^* \right|^2 1_{\{\tau(B)>n\}} = \left| \theta'_{n} - \theta^* \right|^2 1_{\{\tau(B)>n\}} + 2\gamma_{n} \left\{ \theta'_{n} - \theta^*, h(\theta_n) \right\} 1_{\{\tau(B)>n\}} + \gamma_{n} 1_{\{\tau(B)>n\}} r'_n \leq (1 - 2\varepsilon \gamma_{a}) \left| \theta'_{n} - \theta^* \right|^2 1_{\{\tau(B)>n\}} + \gamma_{a} 1_{\{\tau(B)>n\}} \left(r'_n + \left\{ \theta'_{n} - \theta^*, h(\theta_n) - h(\theta'_n) \right\} \right),
\]

where \( r'_n \rightarrow 0 \) almost surely as \( n \rightarrow \infty. \) Since \( \theta_n \) remains in the compact \( \Theta_c, \) \( h \) is continuous and since \( |\theta'_n - \theta| \rightarrow 0, \) it follows that \( \left\{ \theta'_n - \theta^*, h(\theta_n) - h(\theta'_n) \right\} \rightarrow 0. \)

We can summarize the situation like this. Writing \( U'_n = \left| \theta'_n - \theta^* \right|^2 1_{\{\tau(B)>n\}} \), we have
\[
U'_{n+1} \leq (1 - 2\varepsilon \gamma_{a}) U_n + \gamma_{a} r''_n, \quad (6.10)
\]
where \( r''_n \rightarrow 0 \) as \( n \rightarrow \infty. \) This implies that \( U_n \rightarrow 0 \) which, given Lemma 6.6, proves the Proposition.

To see why \( U_n \rightarrow 0, \) let \( \delta > 0 \) be given. Take \( n_0 > 0 \) such that for \( n \geq n_0, \)
\[
\left| r'_n \right| \leq 2\varepsilon \delta \text{ and } (1 - 2\varepsilon \gamma_{a}) 1_{\{\tau(B)>n\}} > 0. \] Then for \( n \geq n_0, (U_{n+1} - \delta) \leq (1 - 2\varepsilon \gamma_{a})(U_n - \delta) + 2\varepsilon \gamma_{a} r''_n/2\varepsilon - \delta \leq (1 - 2\varepsilon \gamma_{a})(U_n - \delta), \)
which implies that \( \limsup(U_n - \delta) \leq 0 \) and, since \( \delta > 0 \) is arbitrary, we conclude that \( \lim U_n = 0. \)
**Proposition 6.2.** Under (A1–3), let $B > 0$ be given. For any function $f \in L_{1/2}^2$, denoting $\tilde{f} = f - \pi^*(f)$, we have:

$$
\frac{1}{n} \sum_{k=1}^{n} \tilde{f}(X_k, I_k) 1_{\{\tau(B) > k-1\}} \to 0, \quad \text{a.s. as } n \to \infty.
$$

**Proof.** In view of Proposition 6.1 and (6.3), we only need show that

$$
\frac{1}{n} \sum_{k=1}^{n} \left( f(X_k, I_k) - \pi_{\theta_k-1}(f) \right) 1_{\{\tau(B) > k-1\}} \to 0 \quad \text{a.s..} (6.12)
$$

Kronecker’s Lemma applied to (6.9) of Lemma 6.6 with $\gamma' = 1/n$, $H_\theta = f$ yields (6.12).

**6.3. Proof of Theorem 4.2**

**Proof.** Assume that [a] hold. Define $\alpha_k = (1 + \gamma_0)^{(1+n_0)k}$ and suppose that

$$
\limsup_{n \to \infty} n(v_n(i) - v_n(j)) = \infty.
$$

This implies the existence of an increasing sequence of integers $\{n_k, \ k \geq 1\}$ such that $n_k(v_n(i) - v_n(j)) > \alpha_k$ and $(X_{n_k}, I_{n_k}) \in \mathcal{X}_i \times \{i\}$, but $(X_{n_k+n_0}, I_{n_k+n_0}) \notin \mathcal{X}_j \times \{j\}$ for all $k \geq 1$. Clearly, $n_k(v_n(i) - v_n(j)) > \alpha_k$ implies that $\theta_{n_k}(i)/\theta_{n_k}(j)$ converges to $+\infty$. But, since $i$ leads to $j$, we can then find $\varepsilon > 0$ and $k_0$ such that, for $k \geq k_0$,

$$
\Pr \left[ (X_{n_k+n_0}, I_{n_k+n_0}) \notin \mathcal{X}_j \times \{j\} \mid \mathcal{F}_{n_k}, (X_{n_k}, I_{n_k}) \in \mathcal{X}_i \times \{i\}, n_k(v_n(i) - v_n(j)) > \alpha_k \right] \leq (1 - \varepsilon).
$$

Thus

$$
\Pr(\limsup_{n \to \infty} n(v_n(i) - v_n(j)) = \infty) \leq \lim_{k \to \infty} (1 - \varepsilon)^k = 0.
$$

Assume that [b] hold. Define $\alpha_k = (1 + \gamma_0)^{2k}$ and suppose that

$$
\limsup_{n \to \infty} n(\max_j v_n(i) - \min_j v_n(j)) = \infty.
$$

Then we can find $i_0 \in \{1, \ldots, d\}$ and an increasing sequence of integers $\{n_k, \ k \geq 1\}$ such that $\min_j v_n(j) = v_n(i_0)$, $n_k(\max_j v_n(j) - v_n(i_0)) > \alpha_k$, $(X_{n_k}, I_{n_k}) \notin \mathcal{X}_0 \times \{i_0\}$ and $(X_{n_k+1}, I_{n_k+1}) \notin \mathcal{X}_0 \times \{i_0\}$ for all $k \geq 1$. Then we can proceed as above and conclude.

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