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Bayesian Statistics and Data Analysis

Lecture 5

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Thanks to Aki Vehtari, Aalto University

- Monte Carlo recap
- Markov Chain Monte Carlo (MCMC)
 - Gibbs sampling
 - Metropolis-Hastings
- MCMC Diagnostics
 - Warm-up
 - Convergence
 - S_{eff} , MCSE, and autocorrelation
- Difficult geometries



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It's all about expectations

$$E_{p(\theta|y)}[f(\theta)] = \int f(\theta) p(\theta|y) d\theta,$$

$$\text{where } p(\theta|y) = \frac{p(y|\theta)p(\theta)}{\int p(y|\theta)p(\theta)d\theta}$$

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We can easily evaluate $p(y|\theta)p(\theta)$ for any θ , but the integral $\int p(y|\theta)p(\theta)d\theta$ is usually difficult.

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- Monte Carlo methods which can sample from $p(\theta^{(s)}|y)$ using only $q(\theta^{(s)}|y)$

$$E_{p(\theta|y)}[f(\theta)] \approx \frac{1}{S} \sum_{s=1}^S f(\theta^{(s)})$$

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- Monte Carlo methods we have discussed so far
 - Inverse CDF works mainly for 1D



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 - Laplace, Variational*, EP* (Ch 4, 13*, next course)



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Markov chains

- Andrey Markov proved weak law of large numbers and central limit theorem for certain dependent-random sequences, which were later named **Markov chains**

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- The probability of each event depends only on the previous event (or finite number of previous events)

$$p(\theta_t | \theta_{t-1}, \theta_{t-2}, \dots) = p(\theta_t | \theta_{t-1})$$

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- Under some assumptions $p(\theta_t | \theta_{t-1})$ will converge (in total variation) to *one* **stationary distribution** $p(\theta)$



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- **Goal in MCMC**: Construct a **transition distribution** with $p(\theta|y)$ as the **stationary distribution**



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- Produce draws $\theta_{(t)}$ given $\theta_{(t-1)}$ from a Markov chain, with stationary distribution $p(\theta|y)$



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 - + combine sequence of easier Monte Carlo draws to form a Markov chain



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 - + combine sequence of easier Monte Carlo draws to form a Markov chain
 - + converges (in total variation) to $p(\theta|y)$ as $t \rightarrow \infty$
 - + central limit theorem holds for expectations
 - draws are dependent
 - construction of an efficient Markov chains is not always easy



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- Random variables $\theta_1, \theta_2, \dots$ where θ_t depends only on the previous $\theta_{(t-1)}$

$$p(\theta_t | \theta_1, \dots, \theta_{(t-1)}) = p(\theta_t | \theta_{(t-1)})$$



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- Choose a transition distribution so the **stationary distribution** of the Markov chain is $p(\theta|y)$



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Subsection 1

Gibbs sampling



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- Alternate sampling from conditional distributions
- Basic algorithm, for $j \in \{1, \dots, J\}$

sample $\theta_{j,t}$ from $p(\theta_j | \theta_{-j,t-1}, y)$,
where $\theta_{-j,t-1} = (\theta_{1,t}, \dots, \theta_{j-1,t}, \theta_{j+1,t-1}, \dots, \theta_{J,t-1})$

- Will converge (in total variation) to $p(\theta|y)$ as $N \rightarrow \infty$



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- j can be multiple (blocked) parameters
- 1D sampling ($|j| = 1$) is generally easy



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- Will converge (in total variation) to $p(\theta|y)$ as $N \rightarrow \infty$
- j can be multiple (blocked) parameters
- 1D sampling ($|j| = 1$) is generally easy
- Popular for **discrete parameters**



Example: Bivariate Normal using Gibbs

$$p(\theta | y) \sim \mathcal{N}(\bar{y}, \frac{1}{n}\Sigma).$$

where Σ is known and $\theta \in \mathcal{R}^2$

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where Σ is known and $\theta \in \mathcal{R}^2$

Using **Gibbs sampling**:

Given starting values $(\theta_1^{(0)}, \theta_2^{(0)})$, for $t = 1, 2, \dots$:

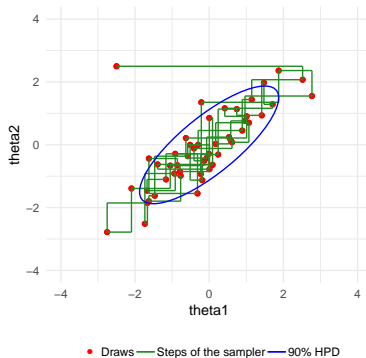
$$\theta_1^{(t)} \sim \mathcal{N}\left(\bar{y}_1 + \frac{\rho}{\sigma_2}(\theta_2^{(t-1)} - \bar{y}_2), \frac{1}{n}\left(\sigma_1 - \frac{\rho^2}{\sigma_2}\right)\right),$$

$$\theta_2^{(t)} \sim \mathcal{N}\left(\bar{y}_2 + \frac{\rho}{\sigma_1}(\theta_1^{(t)} - \bar{y}_1), \frac{1}{n}\left(\sigma_2 - \frac{\rho^2}{\sigma_1}\right)\right).$$

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demo



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 - BUGS / WinBUGS / OpenBUGS / JAGS



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 - For not so easy conditionals, use e.g. inverse-CDF



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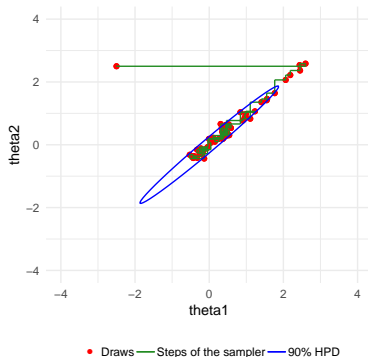


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- Several parameters can be updated in blocks (*blocking*)
- Slow if parameters are highly dependent in the posterior...



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Sampling conditional vs joint

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- How about sampling θ jointly?
 - e.g. it is easy to sample from multivariate normal



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- How about sampling θ jointly?
 - e.g. it is easy to sample from multivariate normal
 - Can we use that to form a Markov chain?



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Subsection 2

Metropolis-Hastings



- Algorithm

1. starting point θ^0

2. $t = 1, 2, \dots$

- (a) pick a proposal θ^* from a **proposal distribution** $J_t(\theta^*|\theta_{t-1})$.

(Proposal distribution has to be symmetric, i.e.

$J_t(\theta_a|\theta_b) = J_t(\theta_b|\theta_a)$, for all θ_a, θ_b)

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$$r = \frac{p(\theta^*|y)}{p(\theta_{t-1}|y)}$$

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(c) set

$$r = \frac{p(\theta^*|y)}{p(\theta_{t-1}|y)}$$
$$\theta_t = \begin{cases} \theta^* & \text{with probability } \min(r, 1) \\ \theta_{t-1} & \text{otherwise} \end{cases}$$



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$$\theta_t = \begin{cases} \theta^* & \text{with probability } \min(r, 1) \\ \theta_{t-1} & \text{otherwise} \end{cases}$$

ie, if $p(\theta^*|y) > p(\theta_{t-1}|y)$ accept the proposal always and otherwise accept the proposal with probability r



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- rejection of a proposal increments the time t also by one ie, the new state is the same as previous



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The Metropolis algorithm

- Algorithm

1. starting point θ^0
2. $t = 1, 2, \dots$
 - (a) pick a proposal θ^* from a **proposal distribution** $J_t(\theta^*|\theta_{t-1})$.
(Proposal distribution has to be symmetric, i.e. $J_t(\theta_a|\theta_b) = J_t(\theta_b|\theta_a)$, for all θ_a, θ_b)
 - (b) calculate acceptance ratio

(c) set

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- rejection of a proposal increments the time t also by one ie, the new state is the same as previous
- step c is executed by generating a random number from $\mathcal{U}(0, 1)$
- $p(\theta^*|y)$ and $p(\theta_{t-1}|y)$ have the same normalization terms, and thus instead of $p(\cdot|y)$, unnormalized $q(\cdot|y)$ can be used, **as the normalization terms cancel out!**



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- Example: one bivariate observation (y_1, y_2)
 - bivariate normal distribution with unknown mean and known covariance

$$\begin{pmatrix} \theta_1 \\ \theta_2 \end{pmatrix} \Big| y \sim \mathcal{N} \left(\begin{pmatrix} y_1 \\ y_2 \end{pmatrix}, \begin{pmatrix} 1 & \rho \\ \rho & 1 \end{pmatrix} \right)$$

- proposal distribution $J_t(\theta^* | \theta_{t-1}) = \mathcal{N}(\theta^* | \theta_{t-1}, \sigma_p^2)$

demo



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- Intuitively more draws from the higher density areas as jumps to higher density are always accepted and only some of the jumps to the lower density are accepted



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- Intuitively more draws from the higher density areas as jumps to higher density are always accepted and only some of the jumps to the lower density are accepted
 - Theoretically
 1. Prove that simulated series is a Markov chain which has unique stationary distribution
 2. Prove that this stationary distribution is the desired target distribution $p(\theta|y)$



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1. Prove that simulated series is a Markov chain which has unique stationary distribution
 - a) irreducible
 - b) aperiodic
 - c) recurrent / not transient



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1. Prove that simulated series is a Markov chain which has unique stationary distribution
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 - = aperiodic (return times are not periodic)
 - holds for a random walk on any proper distribution (except for trivial exceptions)
 - c) recurrent / not transient
 - = probability to return to a state i is 1 as $T \rightarrow \infty$
 - holds for a random walk on any proper distribution (except for trivial exceptions)



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Why Metropolis algorithm works

2. Prove that this stationary distribution is the desired target distribution $p(\theta|y)$: see book

- Show *detailed balance* with respect to $p(\theta|y)$:

$$p(\theta|y) T(\theta|\theta') = p(\theta'|y) T(\theta'|\theta), \quad \forall \theta, \theta'.$$

Here, $T(\theta'|\theta)$ denotes the one-step transition probability (or density).



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More indepth in a proper Markov Theory course!



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- Generalization of Metropolis algorithm for **non-symmetric** proposal distributions
 - acceptance ratio includes ratio of proposal distributions

$$r = \frac{p(\theta^*|y)/J_t(\theta^*|\theta_{t-1})}{p(\theta_{t-1}|y)/J_t(\theta_{t-1}|\theta^*)}$$



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Metropolis-Hastings algorithm

- Ideal proposal distribution is the distribution itself
 - $J(\theta^*|\theta) \equiv p(\theta^*|y)$ for all θ
 - acceptance probability is 1
 - independent draws
 - not usually feasible

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- After the proposal distribution shape has been selected, it is important to select the scale
 - small scale
 - many steps accepted, but the chain moves slowly due to small steps
 - big scale
 - long steps proposed, but many of those rejected and again chain moves slowly

demo



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demo

- Generic rule for rejection rate is 60-90%

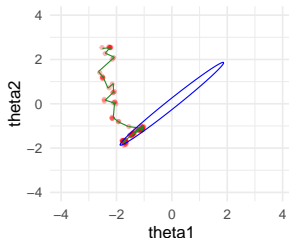


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- Specific case of Metropolis-Hastings algorithm
 - single updated (or blocked)
 - proposal distribution is the conditional distribution
 - proposal and target distributions are same
 - acceptance probability is 1

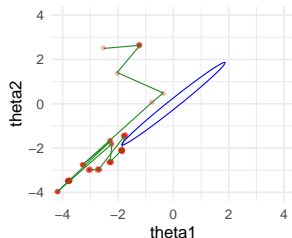


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- Usually doesn't scale well to high dimensions
 - if the shape doesn't match the whole distribution, the efficiency drops



• Draws — Steps of the sampler — 90% HPI



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Section 3

MCMC Diagnostics



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What we want to know

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1. Has the chains **converged** to the posterior distribution?



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What we want to know

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1. Has the chains **converged** to the posterior distribution?
2. How many **efficient** samples do I have?



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Subsection 1

Warm-up



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Warm-up

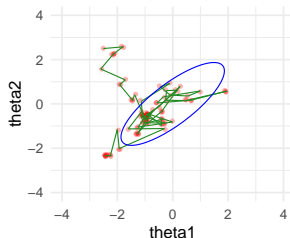
- Asymptotically chain spends the $\alpha\%$ of time where $\alpha\%$ posterior mass is

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Warm-up

- Asymptotically chain spends the $\alpha\%$ of time where $\alpha\%$ posterior mass is
 - but in finite time the initial part of the Markov chain may be non-representative



• Draws — Steps of the sampler — 90% HPD

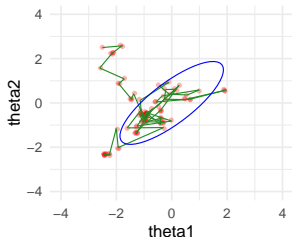
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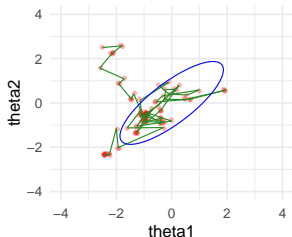
- Warm-up period = (non-representative) draws from the beginning of the Markov chain
 - remove warm-up before using samples for inference
 - warm-up may include phase for adapting algorithm parameters



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- Warm-up period = (non-representative) draws from the beginning of the Markov chain
 - remove warm-up before using samples for inference
 - warm-up may include phase for adapting algorithm parameters
- Also called **burn-in**



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Subsection 2

Convergence



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Assesing convergence

- Several Markov chains make convergence diagnostics easier

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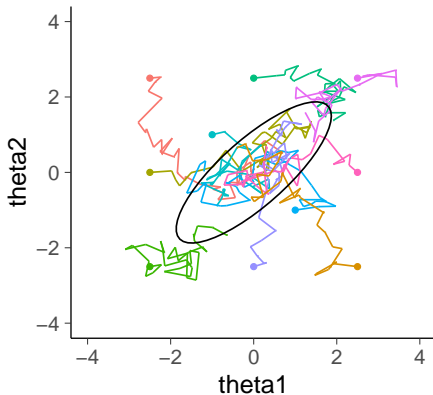


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Assesing convergence

- Several Markov chains make convergence diagnostics easier
- Start chains from different starting points – preferably overdispersed

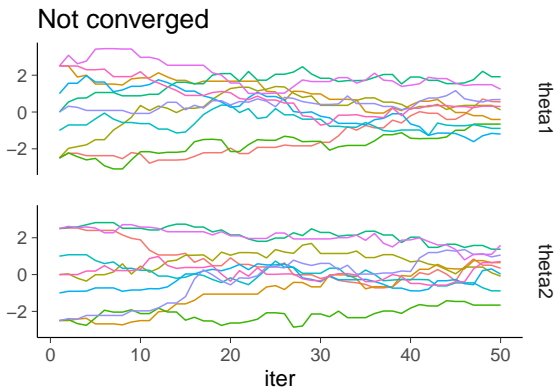
No convergence



- Remove warm-up draws and run chains long enough

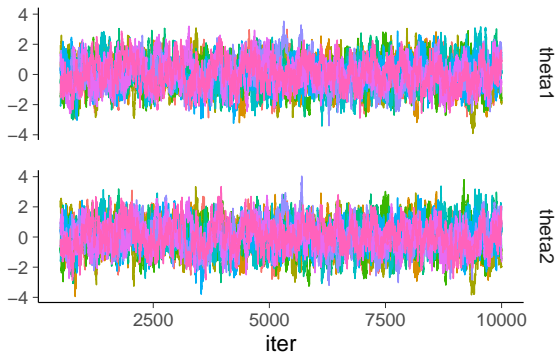


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Visually converged

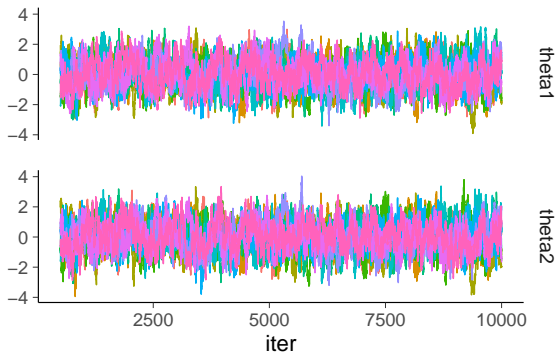


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Visually converged



Visual convergence check is not sufficient



\hat{R} : comparison of within and between variances of the chains

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- \hat{R} or **potential scale reduction factor** (PSRF)
- Compare means and variances of the chains

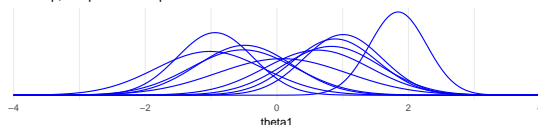


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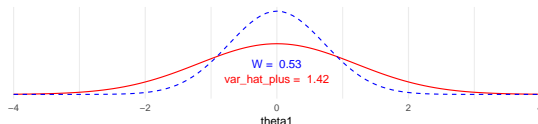
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- \hat{R} or **potential scale reduction factor** (PSRF)
- Compare means and variances of the chains
 W = within chain variance estimate
 var_hat_plus = total variance estimate

50 warmup, 50 post warmup iterations



Rhat = 1.64



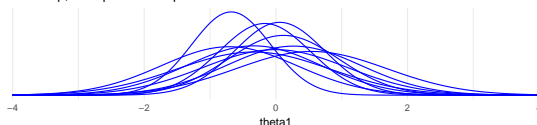


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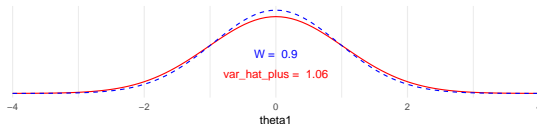
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500 warmup, 500 post warmup iterations



Rhat = 1.08

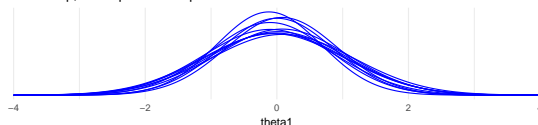




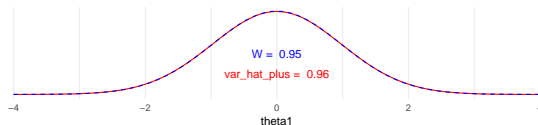
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Rhat = 1





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\hat{R}

- M chains, each having N draws

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- M chains, each having N draws
- Within chains variance W

$$W = \frac{1}{M} \sum_{m=1}^M s_m^2, \text{ where } s_m^2 = \frac{1}{N-1} \sum_{n=1}^N (\theta_{nm} - \bar{\theta}_{.m})^2$$



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- Between chains variance B

$$B = \frac{N}{M-1} \sum_{m=1}^M (\bar{\theta}_{.m} - \bar{\theta}_{..})^2,$$

$$\text{where } \bar{\theta}_{.m} = \frac{1}{N} \sum_{n=1}^N \theta_{nm}, \bar{\theta}_{..} = \frac{1}{M} \sum_{m=1}^M \bar{\theta}_{.m}$$



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 - S_{eff} , MCSE, and autocorrelation
- Difficult geometries

- M chains, each having N draws
- Within chains variance W

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- B/N is variance of the means of the chains



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- B/N is variance of the means of the chains
- Estimate total variance $\text{var}(\theta|y)$ as a weighted mean of W and B

$$\widehat{\text{var}}^+(\theta|y) = \frac{N-1}{N} W + \frac{1}{N} B$$



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- this *overestimates* marginal posterior variance if the starting points are overdispersed



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- Given finite N , W *underestimates* marginal posterior variance
 - single chains have not yet visited all points in the distribution
 - when $N \rightarrow \infty$, $E(W) \rightarrow \text{var}(\theta|y)$



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 - when $N \rightarrow \infty$, $E(W) \rightarrow \text{var}(\theta|y)$
- As $\widehat{\text{var}}^+(\theta|y)$ overestimates and W underestimates, compute

$$\hat{R} = \sqrt{\frac{\widehat{\text{var}}^+}{W}}$$



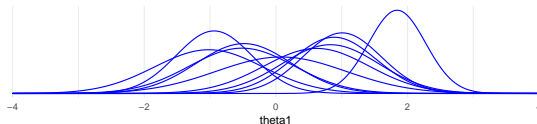
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\hat{R}

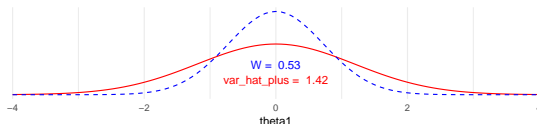
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- BDA3: \hat{R} aka *potential scale reduction factor* (PSRF)
- Compare means and variances of the chains
W = within chain variance estimate
var_hat_plus = total variance estimate

50 warmup, 50 post warmup iterations



Rhat = 1.64

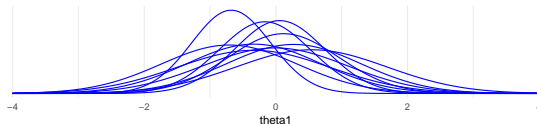




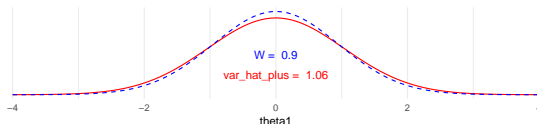
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500 warmup, 500 post warmup iterations



Rhat = 1.08





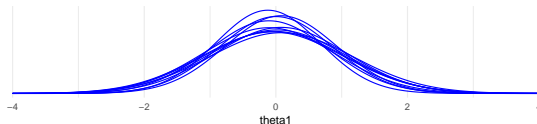
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\hat{R}

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Rhat = 1





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\hat{R}

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$$\hat{R} = \sqrt{\frac{\widehat{\text{var}}^+}{W}}$$

- Estimates how much the scale could reduce if $N \rightarrow \infty$
- $\hat{R} \rightarrow 1$, when $N \rightarrow \infty$
- If \hat{R} is big (e.g., $R > 1.01$), keep sampling



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\hat{R}

$$\hat{R} = \sqrt{\frac{\widehat{\text{var}}^+}{W}}$$

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- Estimates how much the scale could reduce if $N \rightarrow \infty$
- $\hat{R} \rightarrow 1$, when $N \rightarrow \infty$
- If \hat{R} is big (e.g., $R > 1.01$), keep sampling
- If \hat{R} close to 1, it is still possible that chains have not converged
 - if starting points were not overdispersed
 - distribution far from normal (especially if infinite variance)
 - just by chance when N is finite



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\hat{R}

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 - **Convergence**
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- Additional \hat{R} methods to assess convergence
 - Split- \hat{R}
 - Examines *mixing* and *stationarity* of chains
 - To examine stationarity chains are split to two parts: compare means and variances of the split chains



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 - Split- \hat{R}
 - Examines *mixing* and *stationarity* of chains
 - To examine stationarity chains are split to two parts: compare means and variances of the split chains
 - Rank normalized \hat{R}
 - Does not requires that the target distribution has finite mean and variance



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- Monte Carlo estimates still valid (central limit theorem)

$$E_{p(\theta|y)}[f(\theta)] \approx \frac{1}{S} \sum_{s=1}^S f(\theta^{(s)})$$

- Estimation of Monte Carlo error is more difficult
 - evaluation of *effective* sample size, S_{eff} .

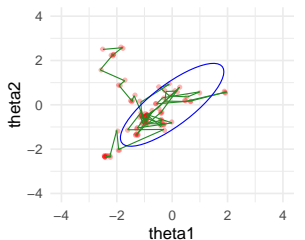


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- Difficult geometries
- Auto correlation function
 - describes the correlation given a certain lag
 - can be used to compare efficiency of MCMC methods



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Autocorrelation

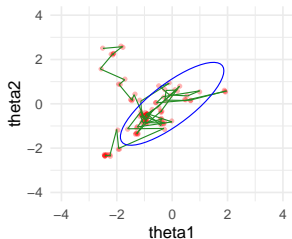


• Draws — Steps of the sampler — 90% HPD

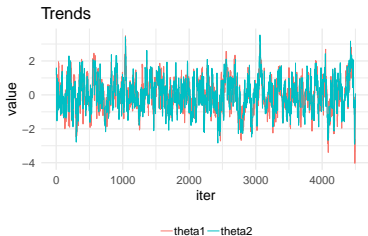


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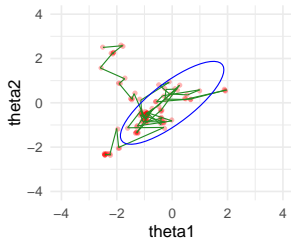




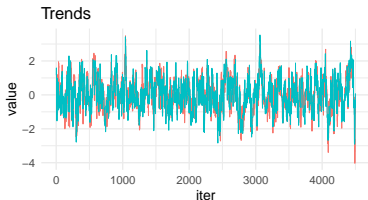
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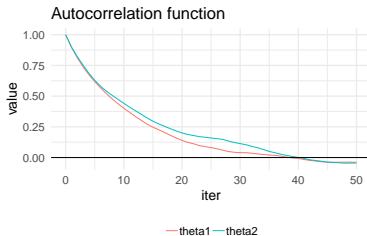
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— theta1 — theta2



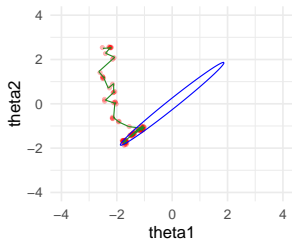
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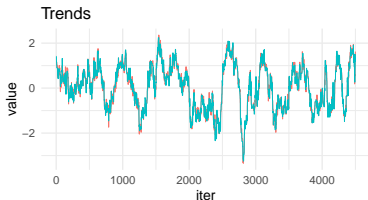
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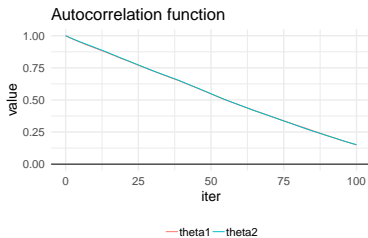
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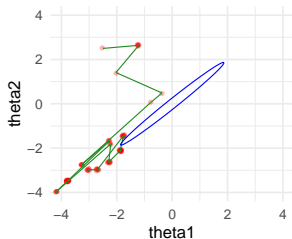
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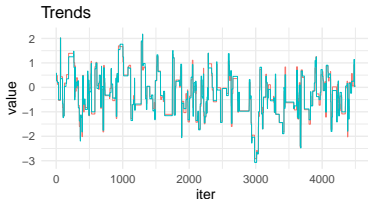
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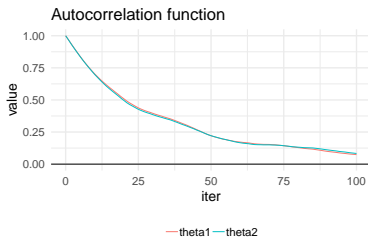
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— θ_1 — θ_2



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- The autocorrelation can be used to estimate Monte Carlo error in case of MCMC
- For expectation $\bar{\theta}$

$$\text{Var}[\bar{\theta}] = \frac{\sigma_{\theta}^2}{S_{\text{eff}}}$$

where $S_{\text{eff}} = S/\tau$, and τ is sum of autocorrelations



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- τ describes how many dependent draws correspond to one independent draw
- Here $S = NM$ (in BDA3 $N = nm$ and $n_{\text{eff}} = N/\tau$)
- BDA3 focuses on S_{eff} and not the Monte Carlo error directly



- Estimation of the autocorrelation using several chains

$$\hat{\rho}_n = 1 - \frac{W - \frac{1}{M} \sum_{m=1}^M \hat{\rho}_{n,m}}{2\widehat{\text{var}}^+}$$

where $\hat{\rho}_{n,m}$ is autocorrelation at lag n for chain m

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The above equation is used in Stan 2.18+



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where $\hat{\rho}_{n,m}$ is autocorrelation at lag n for chain m

- BDA3 has slightly different and less accurate equation. The above equation is used in Stan 2.18+
- Compared to a method which computes the autocorrelation from a single chain, the multi-chain estimate has smaller variance



- Estimation of τ

$$\tau = 1 + 2 \sum_{t=1}^{\infty} \hat{\rho}_t$$

where $\hat{\rho}_t$ is empirical autocorrelation

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- noise is larger for longer lags (less observations)

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- empirical autocorrelation function is noisy
- noise is larger for longer lags (less observations)
- less noisy estimate is obtained by truncating

$$\hat{\tau} = 1 + 2 \sum_{t=1}^T \hat{\rho}_t$$

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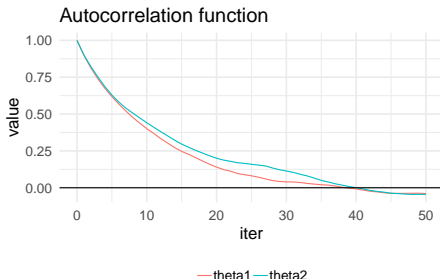
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- Truncation T can be decided adaptively
 - for stationary, irreducible, recurrent Markov chain
 - let $\Gamma_m = \rho_{2m} + \rho_{2m+1}$, which is sum of two consequent autocorrelations
 - Γ_m is positive, decreasing and convex function of m



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Geyer's adaptive window estimator of τ

- Truncation T can be decided adaptively
 - for stationary, irreducible, recurrent Markov chain
 - let $\Gamma_m = \rho_{2m} + \rho_{2m+1}$, which is sum of two consequent autocorrelations
 - Γ_m is positive, decreasing and convex function of m
- Initial positive sequence estimator (Geyer's IPSE)
 - Choose the largest m so, that all values of the sequence $\hat{\Gamma}_1, \dots, \hat{\Gamma}_m$ are positive





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Effective sample size, S_{eff}

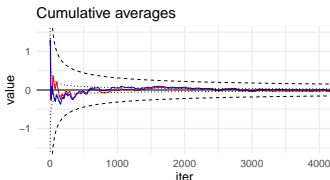
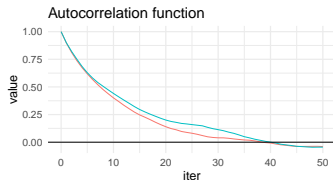
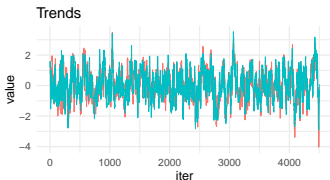
Effective sample size $\text{ESS} = S_{\text{eff}} \approx S/\hat{\tau}$

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Effective sample size, S_{eff}

Effective sample size $\text{ESS} = S_{\text{eff}} \approx S/\hat{\tau}$



-theta1 -theta2 - 95% interval for MCMC error...95% interval for indepen

$$\hat{\tau} = 1 + 2 \sum_{t=1}^T \hat{\rho}_t$$

$$\approx 24$$

$$\hat{S}_{\text{eff}} = \frac{4500}{24} \approx 188$$

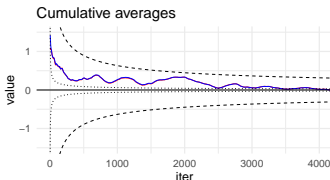
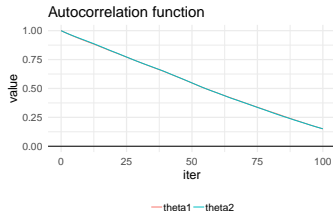
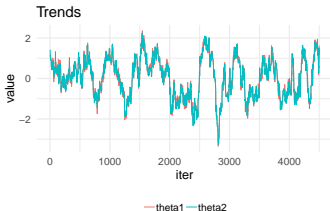
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$$\text{Effective sample size ESS} = S_{\text{eff}} \approx S / \hat{\tau}$$



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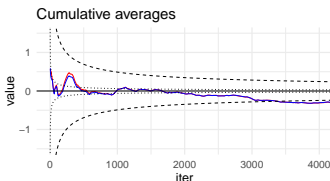
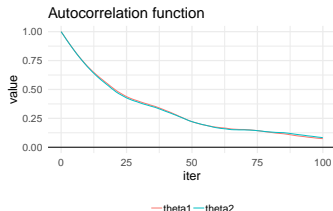
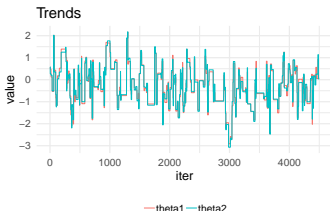
$$\approx 104$$

$$\hat{S}_{\text{eff}} = \frac{4500}{104} \approx 43$$



Effective sample size, S_{eff}

Effective sample size $\text{ESS} = S_{\text{eff}} \approx S/\hat{\tau}$



— theta1 — theta2 - - 95% interval for MCMC error ··· 95% interval for indepen

$$\hat{\tau} = 1 + 2 \sum_{t=1}^T \hat{\rho}_t$$

$$\approx 63$$

$$\hat{S}_{\text{eff}} = \frac{4500}{63} \approx 71$$



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- Difficult geometries

Section 4

Difficult geometries



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Problematic distributions

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demo



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demo
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demo
- Long-tailed with non-finite variance and mean
 - central limit theorem for expectations does not hold