Probabilistic programming for birth-death models of evolution using an alive particle filter with delayed sampling

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SUPPLEMENTARY MATERIAL

A PROOF OF THE UNBIASEDNESS OF THE MARGINAL LIKELIHOOD ESTIMATOR OF THE EXTENDED APF

In this section we prove that the marginal likelihood estimator

$$\widehat{Z} = \prod_{t=1}^{T} \frac{\sum\limits_{n=1}^{N} w_t^{(n)}}{P_t - 1},$$

produced by the extended alive particle filter (APF) for the state space model (Figure 1)

$$\begin{aligned} &x_0 \sim p(x_0), \\ &x_t \sim f_t(x_t|x_{t-1}), \text{ for } t = 1, 2, \ldots, T, \\ &y_t \sim g_t(y_t|x_t), \end{aligned}$$

is unbiased in the sense that $\mathbb{E}[\widehat{Z}] = p(y_{1:T})$.

The structure of our proof is similar to that of Pitt et al. (2012) for the Auxiliary Particle Filter. Let $\mathcal{F}_t = \{x_t^{(n)}, w_t^{(n)}\}_{n=1}^N$ denote the internal state of the particle filter, i.e., the states and weights of all particles, at time t.

Lemma 1.

$$\mathbb{E}\left[\frac{\sum_{n=1}^N w_t^{(n)}}{P_t - 1}\middle|\mathcal{F}_{t-1}\right] = \sum_{n=1}^{N} \frac{w_{t-1}^{(n)}}{\sum_{m=1}^N w_{t-1}^{(m)}} p\left(y_t\middle|x_{t-1}^{(n)}\right).$$

Proof. In the interest of brevity, we will omit conditioning on $\mathcal{F}_{t-1}$ in the notation. For each particle, the APF constructs a candidate sample x' by drawing a sample from $\{x_{t-1}^{(n)}\}$ with the probabilities proportional to the weights $\{w_{t-1}^{(n)}\}$ and propagating it forward to time t such that

$$x' \sim \sum_{n=1}^{N} \frac{w_{t-1}^{(n)}}{\sum_{m=1}^N w_{t-1}^{(m)}} f_t\left(x'\middle|x_{t-1}^{(n)}\right).$$

If $g_t(y_t|x') = 0$, the candidate sample is rejected and the procedure is repeated until acceptance (when $g_t(y_t|x') > 0$).

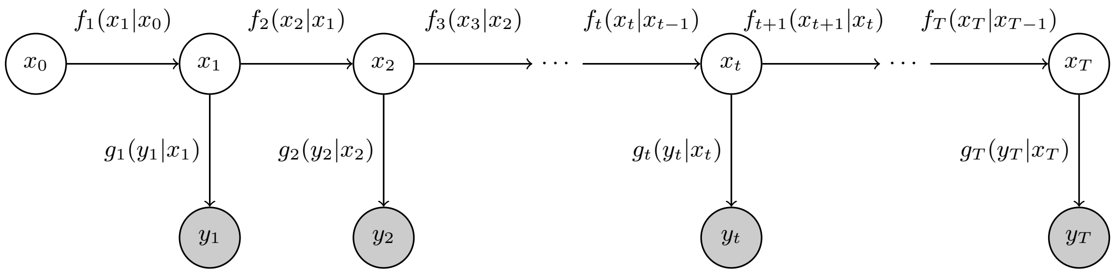

Figure 1: Graphical model of the state space model.

Let $A_t = \{x' : g_t(y_t|x') > 0\}$. The acceptance probability p_{A_t} is then given by

$$p_{A_t} = \int \mathbf{1}_{A_t}(x') \sum_{n=1}^{N} \frac{w_{t-1}^{(n)}}{\sum_{m=1}^{N} w_{t-1}^{(m)}} f_t\left(x' \middle| x_{t-1}^{(n)}\right) dx',$$

where $\mathbf{1}$ denotes the indicator function.

The accepted samples are distributed according to the following distribution:

$$x_t \sim \frac{\mathbf{1}_{A_t}(x_t)}{p_{A_t}} \sum_{n=1}^{N} \frac{w_{t-1}^{(n)}}{\sum_{m=1}^{N} w_{t-1}^{(m)}} f_t\left(x_t \middle| x_{t-1}^{(n)}\right).$$

The expected value of the weight $w_t = g_t(y_t|x_t)$ of an accepted sample is given by

$$\mathbb{E}[w_t] = \int g_t(y_t|x_t) \frac{\mathbf{1}_{A_t}(x_t)}{p_{A_t}} \sum_{n=1}^{N} \frac{w_{t-1}^{(n)}}{\sum_{m=1}^{N} w_{t-1}^{(m)}} f_t\left(x_t \middle| x_{t-1}^{(n)}\right) dx_t.$$

The factor $\mathbf{1}_{A_t}(x_t)$ can be omitted since $\mathbf{1}_{A_t}(x_t) = 0 \Leftrightarrow g_t(y_t|x_t) = 0$, resulting in

$$\begin{aligned}
\mathbb{E}[w_t] &= \int \frac{1}{p_{A_t}} \sum_{n=1}^{N} \frac{w_{t-1}^{(n)}}{\sum_{m=1}^{N} w_{t-1}^{(m)}} f_t\left(x_t \middle| x_{t-1}^{(n)}\right) g_t(y_t|x_t) dx_t \\
&= \frac{1}{p_{A_t}} \sum_{n=1}^{N} \frac{w_{t-1}^{(n)}}{\sum_{m=1}^{N} w_{t-1}^{(m)}} \int f_t\left(x_t \middle| x_{t-1}^{(n)}\right) g_t(y_t|x_t) dx_t \\
&= \frac{1}{p_{A_t}} \sum_{n=1}^{N} \frac{w_{t-1}^{(n)}}{\sum_{m=1}^{N} w_{t-1}^{(m)}}\, p\left(y_t \middle| x_{t-1}^{(n)}\right).
\end{aligned}$$

The APF repeats drawing new samples until $N+1$ samples have been accepted. The total number of draws of candidate samples at time t, P_t, is itself a random variable distributed according to the negative binomial distribution

$$P(P_t = D) = \binom{D-1}{(N+1)-1} p_{A_t}^{N+1} (1 - p_{A_t})^{D-(N+1)}$$

with the support $D \in \{N+1, N+2, N+3, \ldots\}$.

Finally, using the fact that $\mathbb{E}[w_t]$ does not depend on the value of P_t,

$$\begin{aligned}
\mathbb{E}\left[\frac{\sum_{n=1}^N w_t^{(n)}}{P_t - 1}\right] &= \sum_{D=N+1}^{\infty} \frac{N\mathbb{E}[w_t]}{D-1} P(P_t = D) = \sum_{D=N+1}^{\infty} \frac{N\mathbb{E}[w_t]}{D-1} \binom{D-1}{N} p_{A_t}^{N+1} (1-p_{A_t})^{D-(N+1)} \\
&= N\mathbb{E}[w_t] \sum_{D=N+1}^{\infty} \frac{1}{D-1} \binom{D-1}{N} p_{A_t}^{N+1} (1-p_{A_t})^{D-(N+1)} \\
&= N\mathbb{E}[w_t] \sum_{D=N+1}^{\infty} \frac{1}{D-1} \frac{(D-1)(D-2)!}{N(N-1)!(D-(N+1))!} p_{A_t}^{N+1} (1-p_{A_t})^{D-(N+1)} \\
&= \mathbb{E}[w_t] p_{A_t}^{N+1} \sum_{D=N+1}^{\infty} \binom{D-2}{D-(N+1)} (1-p_{A_t})^{D-(N+1)} \\
&\text{(using the binomial theorem)} \\
&= \mathbb{E}[w_t] p_{A_t}^{N+1} p_{A_t}^{-N} = \frac{1}{p_{A_t}} \sum_{n=1}^N \frac{w_{t-1}^{(n)}}{\sum_{m=1}^N w_{t-1}^{(m)}}\, p\left(y_t \middle| x_{t-1}^{(n)}\right) p_{A_t} \\
&= \sum_{n=1}^N \frac{w_{t-1}^{(n)}}{\sum_{m=1}^N w_{t-1}^{(m)}}\, p\left(y_t \middle| x_{t-1}^{(n)}\right).
\end{aligned}$$

$\square$

Lemma 2.

$$\mathbb{E}\left[\frac{\sum_{n=1}^N w_t^{(n)} p\left(y_{t+1:t'} \middle| x_t^{(n)}\right)}{P_t - 1} \middle| \mathcal{F}_{t-1}\right] = \sum_{n=1}^N \frac{w_{t-1}^{(n)}}{\sum_{m=1}^N w_{t-1}^{(m)}} p\left(y_{t:t'} \middle| x_{t-1}^{(n)}\right).$$

Proof. Similar to the proof of Lemma 1 we have that

$$\begin{aligned}
\mathbb{E}[w_t p(y_{t+1:t'}|x_t)] &= \int \frac{1}{p_{A_t}} \sum_{n=1}^N \frac{w_{t-1}^{(n)}}{\sum_{m=1}^N w_{t-1}^{(m)}} f_t\left(x_t \middle| x_{t-1}^{(n)}\right) g_t(y_t|x_t) p(y_{t+1:t'}|x_t) dx_t \\
&= \frac{1}{p_{A_t}} \sum_{n=1}^N \frac{w_{t-1}^{(n)}}{\sum_{m=1}^N w_{t-1}^{(m)}} \int f_t\left(x_t \middle| x_{t-1}^{(n)}\right) g_t(y_t|x_t) p(y_{t+1:t'}|x_t) dx_t \\
&= \frac{1}{p_{A_t}} \sum_{n=1}^N \frac{w_{t-1}^{(n)}}{\sum_{m=1}^N w_{t-1}^{(m)}}\, p\left(y_{t:t'} \middle| x_{t-1}^{(n)}\right)
\end{aligned}$$

and using this result we have that

$$\begin{aligned}
\mathbb{E}\left[\frac{\sum_{n=1}^N w_t^{(n)} p\left(y_{t+1:t'} \middle| x_t^{(n)}\right)}{P_t - 1}\right] &= \sum_{D=N+1}^{\infty} \frac{N\mathbb{E}[w_t p(y_{t+1:t'}|x_t)]}{D-1} \binom{D-1}{N} p_{A_t}^{N+1} (1-p_{A_t})^{D-(N+1)} \\
&= N\mathbb{E}[w_t p(y_{t+1:t'}|x_t)] \sum_{D=N+1}^{\infty} \frac{1}{D-1} \binom{D-1}{N} p_{A_t}^{N+1} (1-p_{A_t})^{D-(N+1)} \\
&= N\mathbb{E}[w_t p(y_{t+1:t'}|x_t)] \frac{p_{A_t}}{N} = N \frac{1}{p_{A_t}} \sum_{n=1}^N \frac{w_{t-1}^{(n)}}{\sum_{m=1}^N w_{t-1}^{(m)}}\, p\left(y_{t:t'} \middle| x_{t-1}^{(n)}\right) \frac{p_{A_t}}{N} \\
&= \sum_{n=1}^N \frac{w_{t-1}^{(n)}}{\sum_{m=1}^N w_{t-1}^{(m)}}\, p\left(y_{t:t'} \middle| x_{t-1}^{(n)}\right).
\end{aligned}$$

$\square$

Lemma 3.

$$\mathbb{E}\left[\prod_{t'=t-h}^{t}\frac{\sum_{n=1}^{N} w_{t'}^{(n)}}{P_{t'}-1}\middle|\mathcal{F}_{t-h-1}\right]=\sum_{n=1}^{N}\frac{w_{t-h-1}^{(n)}}{\sum_{m=1}^{N} w_{t-h-1}^{(m)}}p\left(y_{t-h:t}\middle|x_{t-h-1}^{(n)}\right).$$

Proof. By induction.

The base step for $h=0$ was proved in Lemma 1.

In the induction step, let us assume that the equality holds for h and prove it for $h+1$:

$$\mathbb{E}\left[\prod_{t'=t-h-1}^{t}\frac{\sum_{n=1}^{N} w_{t'}^{(n)}}{P_t'-1}\middle|\mathcal{F}_{t-h-2}\right]=\mathbb{E}\left[\mathbb{E}\left[\prod_{t'=t-h}^{t}\frac{\sum_{n=1}^{N} w_{t'}^{(n)}}{P_t'-1}\middle|\mathcal{F}_{t-h-1}\right]\frac{\sum_{n=1}^{N} w_{t-h-1}^{(n)}}{P_{t-h-1}-1}\middle|\mathcal{F}_{t-h-2}\right]$$

(using the induction assumption)

$$=\mathbb{E}\left[\sum_{n=1}^{N}\frac{w_{t-h-1}^{(n)}}{\sum_{m=1}^{N} w_{t-h-1}^{(m)}}p\left(y_{t-h:t}\middle|x_{t-h-1}^{(n)}\right)\frac{\sum_{n=1}^{N} w_{t-h-1}^{(n)}}{P_{t-h-1}-1}\middle|\mathcal{F}_{t-h-2}\right]$$

$$=\mathbb{E}\left[\sum_{n=1}^{N}\frac{w_{t-h-1}^{(n)}}{P_{t-h-1}-1}p\left(y_{t-h:t}\middle|x_{t-h-1}^{(n)}\right)\middle|\mathcal{F}_{t-h-2}\right]$$

(using Lemma 2)

$$=\sum_{n=1}^{N}\frac{w_{t-h-2}^{(n)}}{\sum_{m=1}^{N} w_{t-h-2}^{(m)}}p\left(y_{t-h-1:t}\middle|x_{t-h-2}^{(n)}\right).$$

□

Theorem 1.

$$\mathbb{E}\left[\prod_{t=1}^{T}\frac{\sum\limits_{n=1}^{N} w_t^{(n)}}{P_t-1}\right]=p(y_{1:T}).$$

Proof. Using Lemma 3 with $t=T, h=T-1$ and

$$\mathbb{E}\left[\frac{1}{N}\sum_{n=1}^{N}p\left(y_{1:T}\middle|x_0^{(n)}\right)\right]=p(y_{1:T}).$$

□

B GENERATIVE MODEL FOR CRBD

The pseudocode for generating phylogenetic trees using the CRBD model is listed in Algorithm 1.

C RELEVANT CONJUGACY RELATIONSHIPS

C.1 NEGATIVE BINOMIAL AND LOMAX DISTRIBUTION

Negative binomial distribution

Parameters: number of successes $k>0$ before the experiment is stopped, probability of success $p\in(0,1)$

Algorithm 1 Pseudocode for generating trees using the CRBD model.

```
function CRBD(τ_orig)
    return (τ_orig, {CRBD'(τ_orig)})

function CRBD'(τ)
    Δ ~ Exponential(λ + μ)
    τ' ← τ − Δ
    if τ' < 0 then
        return (0, ∅)
    e ~ Cat(p_1 = λ/(λ+μ), p_2 = μ/(λ+μ))
    if e = 1 then
        return (τ', {CRBD'(τ'), CRBD'(τ')})
    else
        return (τ', ∅)
```

Probability mass function:

$$f(r|k,p) = \binom{r+k-1}{k-1} p^k (1-p)^r \text{ for } r \in \mathbb{N} \cup \{0\},$$

where r is the number of failures.

Lomax distribution

Parameters: scale $\lambda > 0$, shape $\alpha > 0$

Probability density function:

$$f(\Delta|\lambda,\alpha) = \frac{\alpha}{\lambda}\left(1 + \frac{\Delta}{\lambda}\right)^{-(\alpha+1)} \text{ for } \Delta \geq 0$$

C.2 CONJUGACY RELATIONSHIPS

Gamma-Poisson mixture

Prior distribution: $\nu \sim \mathrm{Gamma}(k,\theta)$ with the probability density function

$$f(\nu|k,\theta) = \frac{1}{\Gamma(k)\theta^k}\nu^{k-1}e^{-\nu/\theta} \text{ for } \nu > 0$$

Likelihood: $n \sim \mathrm{Poisson}(\nu\Delta)$ with the probability mass function

$$f(n|\lambda) = \frac{\lambda^n}{n!}e^{-\lambda} \text{ for } n \in \mathbb{N} \cup \{0\},$$

where $\lambda = \nu\Delta$.

Prior predictive distribution ($k \in \mathbb{N}$):

$$\begin{aligned} f(n|k,\theta) &= \int_0^\infty \frac{1}{\Gamma(k)\theta^k}\nu^{k-1}e^{-\nu/\theta} \times \frac{(\nu\Delta)^n}{n!}e^{-\nu\Delta}d\nu = \frac{\Delta^n}{n!(k-1)!\theta^k}\int_0^\infty \nu^{n+k-1}e^{-\nu(1/\theta+\Delta)}d\nu \\ &= \frac{\Delta^n}{n!(k-1)!\theta^k}\left(\frac{1}{\theta}+\Delta\right)^{-(n+k)}(n+k-1)! = \binom{n+k-1}{k-1}\left(\frac{1}{1+\Delta\theta}\right)^k\left(1-\frac{1}{1+\Delta\theta}\right)^n \end{aligned}$$

$$n|k,\theta \sim \mathrm{NegativeBinomial}\left(k, \frac{1}{1+\Delta\theta}\right)$$

Posterior distribution:

$$f(\nu|n) \propto \frac{1}{\Gamma(k)\theta^k}\nu^{k-1}e^{-\nu/\theta} \times \frac{(\nu\Delta)^n}{n!}e^{-\nu\Delta} \propto \nu^{k+n-1}e^{-\nu(1/\theta+\Delta)} = \nu^{(k+n)-1}e^{-\nu/\left(\frac{\theta}{1+\Delta\theta}\right)}$$

$$\nu|n \sim \text{Gamma}\left(k+n, \frac{\theta}{1+\Delta\theta}\right)$$

Gamma-exponential mixture

Prior distribution: $\nu \sim \text{Gamma}(k, \theta)$

Likelihood: $\Delta \sim \text{Exponential}(\nu)$ with the probablity density function

$$f(\Delta|\nu) = \nu e^{-\nu\Delta} \text{ for } \Delta \geq 0$$

Prior predictive distribution:

$$\begin{aligned} f(\Delta|k,\theta) &= \int_0^\infty \frac{1}{\Gamma(k)\theta^k}\nu^{k-1}e^{-\nu/\theta} \times \nu e^{-\nu\Delta}d\nu = \frac{1}{\Gamma(k)\theta^k}\int_0^\infty \nu^k e^{-\nu(1/\theta+\Delta)}d\nu \\ &= \frac{1}{\Gamma(k)\theta^k}\left(\frac{1}{\theta+\Delta}\right)^{-(k+1)}\Gamma(k+1) = \frac{k}{\theta^k}\left(\frac{1}{\theta}+\Delta\right)^{-(k+1)} = k\theta(1+\Delta\theta)^{-(k+1)} \end{aligned}$$

$$\Delta|k,\theta \sim \text{Lomax}\left(\frac{1}{\theta}, k\right)$$

Posterior distribution:

$$f(\nu|\Delta) \propto \frac{1}{\Gamma(k)\theta^k}\nu^{k-1}e^{-\nu/\theta} \times \nu e^{-\nu\Delta} \propto \nu^k e^{-\nu(1/\theta+\Delta)} = \nu^{(k+1)-1}e^{-\nu/\left(\frac{\theta}{1+\Delta\theta}\right)}$$

$$\nu|\Delta \sim \text{Gamma}\left(k+1, \frac{\theta}{1+\Delta\theta}\right)$$

D SOURCE CODE

Birch is available at

`https://birch-lang.org/`

The source code for the CRBD and BiSSE models is available at

`https://github.com/kudlicka/paper-2019-probabilistic`

References

M. K. Pitt, R. dos Santos Silva, P. Giordani, and R. Kohn. On some properties of Markov chain Monte Carlo simulation methods based on the particle filter. *Journal of Econometrics*, 171(2):134–151, 2012.

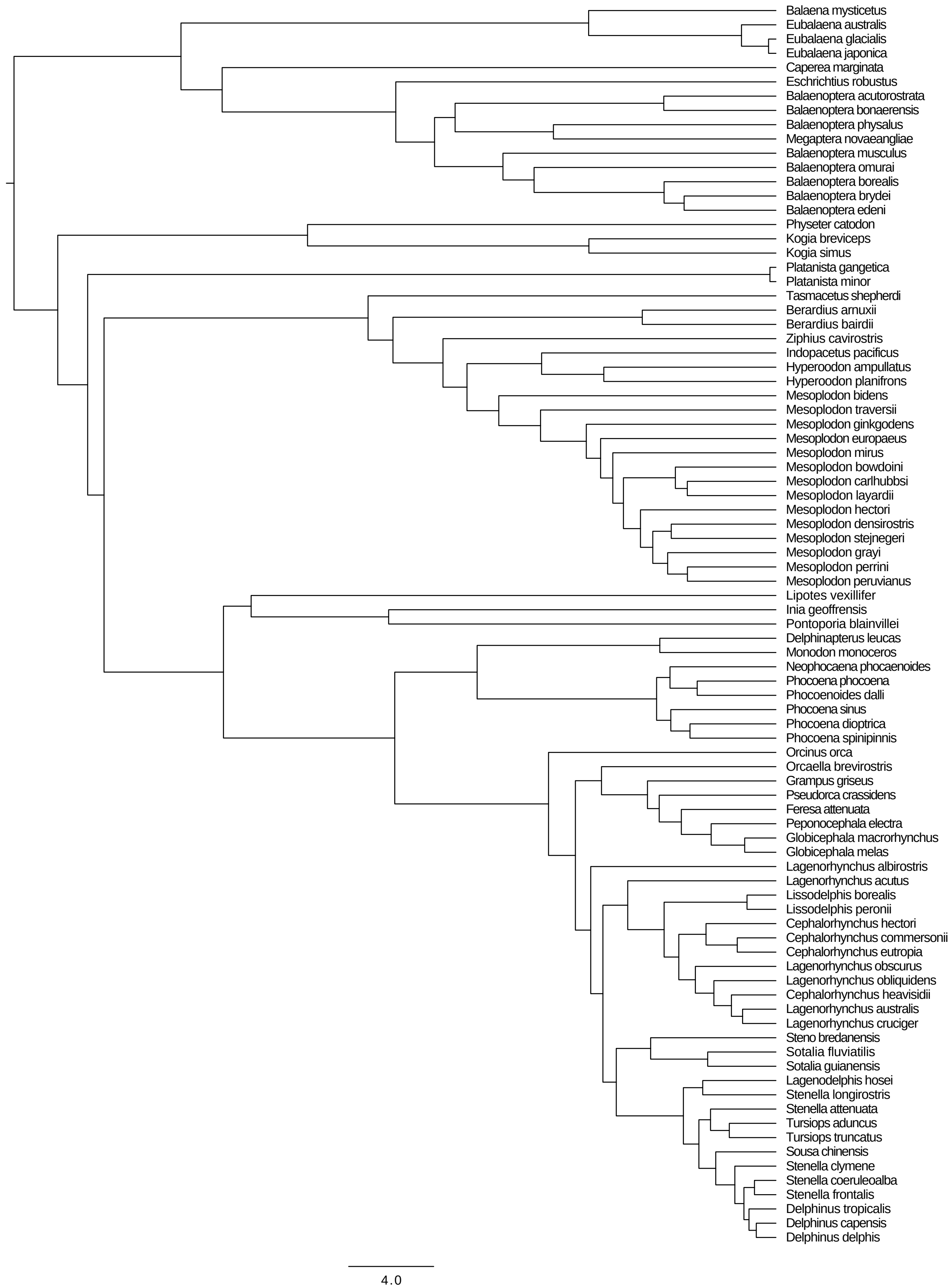

Figure 2: Phylogeny of cetaceans (whales, dolphins and porpoises).