

1 **Quantifying the limits to forecasting the impacts of environmental change on the**
2 **Yellowstone bison population**

3 Andrew T. Tredennick¹

4 ¹*Department of Wildland Resources and the Ecology Center, 5230 Old Main Hill, Utah State University, Logan,*
5 *Utah 84322 USA*

6 **Running Head:** Limits to bison forecasting

7 **Corresponding Author:**

8 Andrew Tredennick

9 Department of Wildland Resources and the Ecology Center

10 Utah State University

11 5230 Old Main Hill

12 Logan, Utah 84322 USA

13 Phone: +1-970-443-1599

14 Fax: +1-435-797-3796

15 Email: atredenn@gmail.com

Abstract

Making ecological forecasts requires bringing together data and models that can project states under future environmental conditions. An interesting, but often overlooked, complication of ecological forecasting is that the state of interest (e.g., population abundance) is often impacted by an environmental covariate (e.g., precipitation) that itself must be forecast. In many cases the limits to ecological forecasting lie in the environmental covariates we choose, because even if those covariates produce better understanding of the system, uncertainty in their forecast will propagate through to impact the uncertainty of the focal state. Thus, understanding when and by how much forecasts are limited by the uncertainty of environmental covariates is an urgent need. Here we use new methods to explicitly quantify the different sources of uncertainty on forecasts of the Yellowstone bison population. We fit a Bayesian state-space model to 38 years of population counts assuming Ricker population growth with an environmental effect of mean annual snow water equivalent. We made forecasts for an additional 10 years and found (1) that including known values of mean annual snow water equivalent increased out-of-sample forecast accuracy and (2) that xx - xx% of forecast error was associated with uncertainty in future projections of snow water equivalent. In comparison, only x% of forecast error was associated with uncertainty in initial conditions (e.g., population state at $t - 1$), and fell to 0% at a forecast horizon of three years. We used simulations to show how forecast error decreased with increased precision of snow water equivalent forecasts. Our findings highlight the necessity of partitioning forecast error into quantifiable components. Only then can key limitations be identified. In our case, more precise observations of bison population abundance will not lead to better forecasts. To better forecast the impacts of environmental change on the Yellowstone bison population, we need better forecasts of the environmental change itself.

Key words: Bayesian state-space model; Bison bison; ecological forecasting; population model; Yellowstone National Park, USA; uncertainty.

Introduction

Forecasting the impacts of environmental change on populations, communities, and ecosystems is a major challenge of the 21st century.

Materials and Methods

Data

Aerial counts of the Yellowstone bison population from 1970-2017 (Fig. 1A) were used to estimate model parameters and states. Counts typically occurred four times a year. We used summer counts only because the summer aggregation of bison tends to produce more accurate counts (Hobbs et al. 2015). Replicate counts were taken in 41 of the 48 years, and from these replicates we estimated the mean annual total population in the park and the annual standard deviation of counts (i.e., observation error).

We used mean annual snow water equivalent (SWE, mm) as an environmental covariate in our population model. SWE data came from the West Yellowstone SNOTEL station (Site #924). We downloaded the SNOTEL data using the `snotelr` package in R and then calculated the mean SWE for each year.

State-space model

We used a Bayesian state-space model to obtain posterior distributions of all unknown parameters, latent states, and forecasts of future states. The fully specified model, including a model of the ecological process, a model linking the process to the data, and parameter models, takes the form:

$$\left[\theta_p, z_{(t)}, z_{(t-1)} | y_{(t)}, x_{(t)}, \sigma_{o,(t)}^2 \right] \propto \underbrace{\prod_{t=2}^{58} \left[z_{(t)} | \theta_p, z_{(t-1)}, x_{(t)} \right]}_{\text{process}} \underbrace{\prod_{t=1}^{48} \left[y_{(t)} | \sigma_{o,(t)}^2, z_{(t)} \right]}_{\text{data}} \underbrace{\left[\theta_p, z_{(t=1)} \right]}_{\text{parameters}}, \quad (1)$$

where the notation $[a|b,c]$ reads as the probability or probability density of a conditional on b and c (Hobbs and Hooten 2015). θ_p is a vector of parameters in the process model, $z_{(t)}$ is the latent, or unobservable and true, state of the bison population abundance at time t , $y_{(t)}$ is the observed state of bison population abundance at time t , and $x_{(t)}$ is the standardized value of mean annual SWE at time t . Note that the product associated with the “process model” applies

over 10 more years than the product associated with the “data model.” The extra ten years are forecasts of the latent state ten years into the future, for which no likelihood can be calculated. Our process model represents the population dynamics of the Yellowstone bison using the stochastic model

$$z_{(t)} \sim \text{lognormal} \left(\log \left(z_{(t-1)} e^{r+b_0 z_{(t-1)}+b_1 x_{(t)}} \right), \sigma_p^2 \right), \quad (2)$$

where the deterministic model $z_{(t-1)} e^{r+b_0 z_{(t-1)}+b_1 x_{(t)}}$ is a Ricker model of population growth (Ricker 19xx) that predicts the mean of $z_{(t)}$ on the log scale as function of the true state of the population at time $t - 1$ ($z_{(t-1)}$), the intrinsic, per capita rate of increase (r), the strength of density-dependence (b_0), and the effect (b_1) of mean annual snow water equivalent at time t ($x_{(t)}$). The quantity σ_p^2 is the process variance on the log scale, which accounts for all the drivers of the true state that are not found in the deterministic model.

We used a normal likelihood to link the observations (y) to the estimated latent states (z). Using the estimates of observation variance for each year t ($\sigma_{o,(t)}^2$), the likelihood is $y_{(t)} \sim \text{Normal} \left(z_{(t)}, \sigma_{o,(t)}^2 \right)$. As mentioned earlier, replicate observations were not available for seven of the 48 years. In those seven years we could not estimate observation error. Therefore, in any year for which observation error could not be estimated, we set observation variance to the maximum value observed across the time series. This is a conservative approach that ensures our estimates of the latent states in those years are not overly confident.

Our state-space model requires prior distributions for all parameters and for the initial condition of z . We used an informative prior for the intrinsic growth rate, r , based on the population growth (λ) reported by Hobbs et al. (2015). The Hobbs et al. (2015) model includes estimates of two population growth rates, one for populations infected with brucellosis and one for populations not infected with brucellosis. Although the population growth rate was depressed in brucellosis-infected populations, the posterior estimate of the difference between the two growth rates overlapped zero, indicating little support for differentiating among the two. We therefore chose to use the estimates for the brucellosis-free population: mean = 1.11 and s.d.

92 = 0.0241. We converted λ , the population growth rate, to r , the per capita rate of increase, by
 93 log-transforming λ : $r = \log(\lambda) = 0.1$. To get the standard deviation on the same scale, we
 94 simulated 100,000 numbers from the distribution $\log(\text{Normal}(1.11, 0.0241))$ and calculated
 95 the standard deviation of those numbers, which equaled 0.02. Thus, our prior distribution
 96 for r was $\text{Normal}(0.1, 0.02)$. We defined the prior distribution of the initial condition $z_{(t=1)}$ as
 97 $z_{(t=1)} \sim \text{Normal}\left(y_{(t=1)}, \sigma_{o,(t=1)}^2\right)$.

98 We chose all other prior distributions to be vague. However, no prior distribution is completely
 99 uninformative, so we made sure that our choice of priors did not have large impacts on posteri-
 100 ors by trying several choices of priors and their associated parameters. We then observed their
 101 effects on the posteriors (Hobbs and Hooten 2015), which were small. Our final chosen priors
 102 were: $1/\sigma_p^2 = \tau_p \sim \text{Gamma}(0.01, 0.01)$ and $b_{\in(0,1)} \sim \text{Normal}(0, 1000)$.

103 We estimated the posterior distributions of parameters and states using Markov chain Monte
 104 Carlo (MCMC) methods implemented using JAGS 4.2.0 (Plummer) and the `rjags` package
 105 (Plummer2) to connect JAGS to R (R citation). Posterior samples were collected from three
 106 chains, each with unique initial values that were variable relative to the posterior distributions.
 107 We used 50,000 iterations to adapt the chains for optimal sampling, then accumulated 100,000
 108 samples from each chain after an initial burn-in period of 50,000 iterations. We checked for
 109 convergence and stationarity of the MCMC chains by visually inspecting the trace plots for each
 110 parameter and state, and by calculating multivariate scale reduction factors and the Gelman
 111 metric (citations) using the `coda` package in R. R code and data necessary to reproduce our
 112 analysis has been archived on Figshare (*link*) and released on GitHub ([https://github.com/](https://github.com/atredennick/bison_forecast/releases)
 113 `atredennick/bison_forecast/releases`).

114 **Model forecasts**

115 **Partitioning forecast uncertainty**

116 To understand the “forecasting problem” in ecology, Dietze (2017) proposed a first-principles
 117 approach to partitioning forecast uncertainty. Consider a dynamical model designed to pre-
 118 dict some state y in the future (y_{t+1}) based on the current state (y_{t+1}), an external covariate (x),

119 parameters (θ), and process error (ϵ). We can then write a general form of the model as:

$$y_{t+1} = f(y_t, x_t | \theta) + \epsilon, \quad (3)$$

120 which states that y at time $t + 1$ is a function of y and x at time t conditional on the model
 121 parameters (θ) plus process error (ϵ). Using a Taylor Expansion, Dietze shows that forecast
 122 variance ($Var[y_{t+1}]$) is:

$$Var[y_{t+1}] = \underbrace{\left(\frac{\delta f}{\delta y}\right)^2}_{\text{stability}} \underbrace{Var[y_t]}_{\text{IC uncert.}} + \underbrace{\left(\frac{\delta f}{\delta x}\right)^2}_{\text{driver sens.}} \underbrace{Var[x_t]}_{\text{driver uncert.}} + \underbrace{\left(\frac{\delta f}{\delta \theta}\right)^2}_{\text{param sens.}} \underbrace{Var[\theta]}_{\text{param. uncert.}} + \underbrace{Var[\epsilon]}_{\text{process error}}, \quad (4)$$

123 where each additive term follows a pattern of *sensitivity* times *variance*. Thus, the variance at-
 124 tributable to any particular factor is a function of how sensitive the model is to the factor and
 125 the variance of that factor. For example, large sensitivity to the covariate x can be compensated
 126 for if the uncertainty of the covariate is low.

127 Results

128 Discussion

129 Acknowledgements

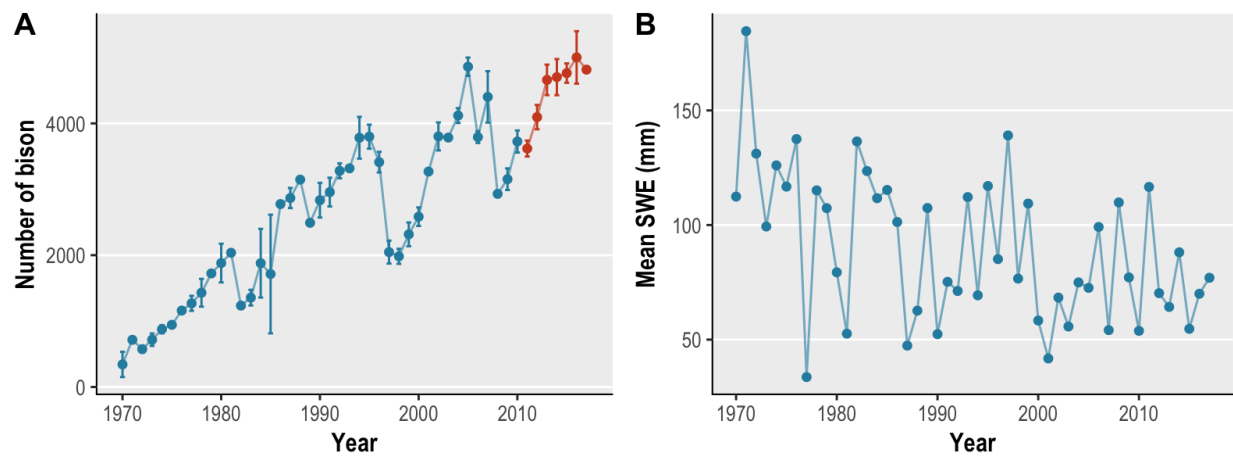


Figure 1: The data.

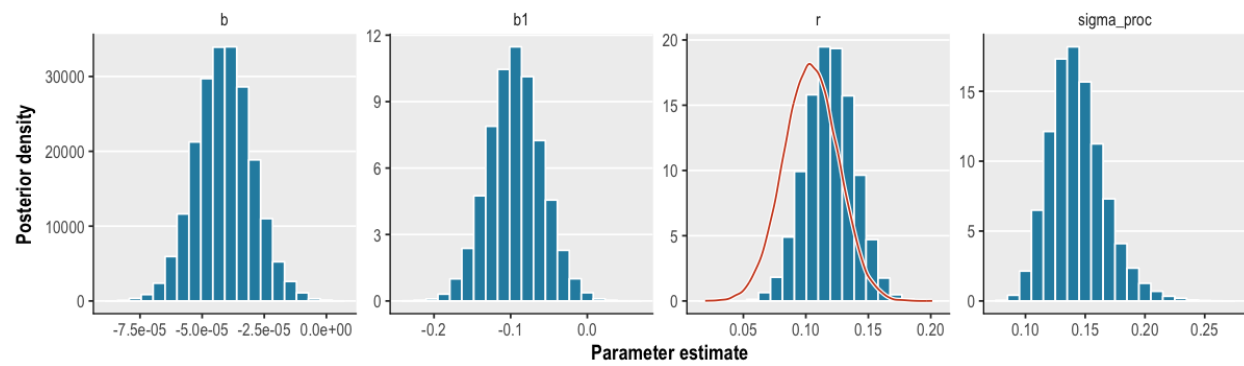


Figure 2: The posteriors.

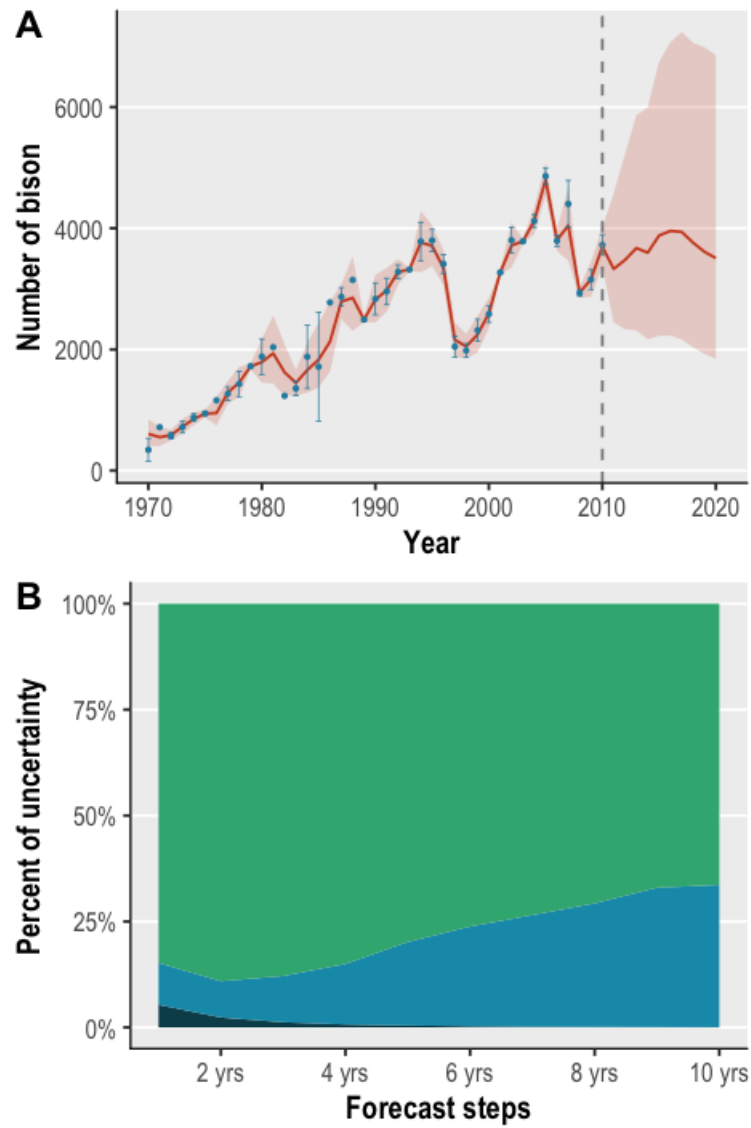


Figure 3: The forecasts.