

Combining Breeding Bird Survey and Christmas Bird Count Data to Evaluate Seasonal Components of Population Change in Northern Bobwhite

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ABSTRACT Annual surveys of wildlife populations provide information about annual rates of change in populations but provide no information about when such changes occur. However, by combining data from 2 annual surveys, conducted in different parts of the year, seasonal components of population change can be estimated. We describe a hierarchical model for simultaneous analysis of 2 continent-scale monitoring programs. The Christmas Bird Count is an early winter survey, whereas the North American Breeding Bird Survey is conducted in June. Combining information from these surveys permits estimation of seasonal population variance components and improves estimation of long-term population trends. The composite analysis also controls for survey-specific sampling effects. We applied the model to estimation of population change in northern bobwhites (*Colinus virginianus*). Over the interval 1969–2004, bobwhite populations declined, with trend estimate of -3.56% per year (95% CI = $[-3.80\%, -3.32\%]$) in the surveyed portion of their range. Our analysis of seasonal population variance components indicated that northern bobwhite populations changed more in the winter and spring portion of the year than in the summer and fall portion of the year. (JOURNAL OF WILDLIFE MANAGEMENT 72(1):44–51; 2008)

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The northern bobwhite (*Colinus virginianus*) has been of conservation concern since early in the 20th century (see review in Williams et al. 2004). Continuing concerns about declines and needs for landscape-scale management to enhance bobwhite populations have led to a variety of conservation activities and research into demography and habitat associations of the species (Guthery et al. 2000, Dimmick et al. 2002, Peterson et al. 2002). Fundamental to the discussion of population change in bobwhite in recent years has been use of information from the North American Breeding Bird Survey (BBS; Sauer et al. 2005a), which provides quantitative information on population change at a variety of geographic scales within the United States portion of bobwhite range. Although the BBS has features that limit its use for some conservation activities (Sauer et al. 2005b), model-based analyses with controls for factors affecting detectability provide a basis for estimation of population change. Such analyses also allow for investigation of relations between population relative abundance or change and environmental features (Link and Sauer 2002).

Hierarchical modeling procedures used in modern BBS analysis can be adapted to allow for simultaneous analysis of BBS data and data from other surveys (Link and Sauer 2002). We present analyses of local and range-wide population change for northern bobwhite based on data from the BBS and the Christmas Bird Count (CBC; Butcher 1990). We used data from across the bobwhite range (comprising 69 strata, defined as the intersection of states and physiographic strata) over the 36-year period from 1969 to 2004. Our analysis took into account the

distinct sampling features of the 2 surveys—observer effects in the BBS and effort effects in the CBC—while modeling features of population change common to the 2 surveys. The temporal staggering of the 2 surveys (BBS, early summer; CBC, early winter) can be exploited to examine seasonal components of annual population change. Bobwhites are nonmigratory, so these components of population change are interpretable in terms of recruitment and mortality effects. The potential value of such an approach is 2-fold: 1) it provides more information for estimation of population change; and 2) it allows estimation of seasonal variation in population change, which can be associated with environmental features influencing change and used as an additional source of change data for comparison with predictions from demographic models.

We begin by describing the 2 surveys and the model we use for analysis of the combined data set. The model includes survey-specific sampling effects and survey-independent population effects. We describe the population effects, and their manifestation in BBS and CBC counts. Next, we describe the survey-specific sampling effects, then the fitting of the model under the Bayesian paradigm, and the construction of composite indices. Our goal in this paper is to communicate the potential and value of the combined analysis while limiting description of technical aspects of the analysis; a more detailed account of a similar analysis applied to Carolina wrens (*Thryothorus ludovicianus*) is given in Link and Sauer (2007). Finally, we present survey-wide summaries and results for 5 selected strata, and compare the composite results to a BBS analysis for the species conducted using the methods of Link and Sauer (2002). To our knowledge, the results presented in this paper

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represent the first analysis of bobwhite data using hierarchical models for estimation of population change and the first composite analysis for the species.

METHODS

North American Breeding Bird Survey

The BBS is a roadside survey that was initiated in 1966 and consists of >4,000 routes in the United States and Canada, of which approximately 3,000 are surveyed annually in early June. Observers conduct 3-minute counts of all birds seen or heard at 50 stops at 0.8-km increments along a 39.5-km route. We based our analysis on the total counts of bobwhites by route. Counts were conducted according to a set of protocols regarding time of day, weather conditions, and observer behavior, with the goal of reducing irrelevant sources of variation in the data. Nevertheless, substantial variation exists among BBS observers (Sauer et al. 1994). There also is evidence of variation within observers (i.e., that the proportion of birds detected by an individual changes through time). Informal analysis of the latter effect indicated that it is limited to and adequately accounted for by a first-year effect, in that observers tend to count fewer bobwhites than expected in their first year. It is crucial that observer effects be included in analysis of BBS data because there is a substantial turnover in observers (median yr of service is roughly 4 or 5), and there is considerable evidence of temporal trend in the proportion of birds counted by new recruits to the BBS (Sauer et al. 1994). See Sauer et al. (2005a) for details of the BBS, its history, data limitations, summaries of analyses for 419 species, and a bibliography.

Christmas Bird Count

Initiated in 1900, the CBC consists of counts made on 24.1-km-diameter circles from mid-December to early January. The CBC is primarily recreational, and there is little standardization of counting methods aside from limiting the counts to those collected by known participants in a single, preselected day. The number of individuals participating and the duration of counts vary among circles and through time (see Butcher 1990 for details of the CBC). In recent years, the number of party-hours has been recorded as a covariate to account for the variable duration of and participation in the count. We refer to this covariate as effort. In Fig. 1, we present a graph of total counts of bobwhite, with and without controls for effort; the reason for the discrepancy in the apparent pattern of population change becomes evident on examining Fig. 2, which displays a simple effort adjustment (mean effort/effort in yr i) as a function of year. The importance of controlling for effort is manifest.

Model

We model counts using an overdispersed Poisson loglinear regression. We assumed that 1) conditional on their means, counts were Poisson random variables; and 2) each mean λ could be decomposed as

$$\log(\lambda) = \text{Sampling Effects} + \text{Population Effects} + \text{Noise.} \quad (1)$$

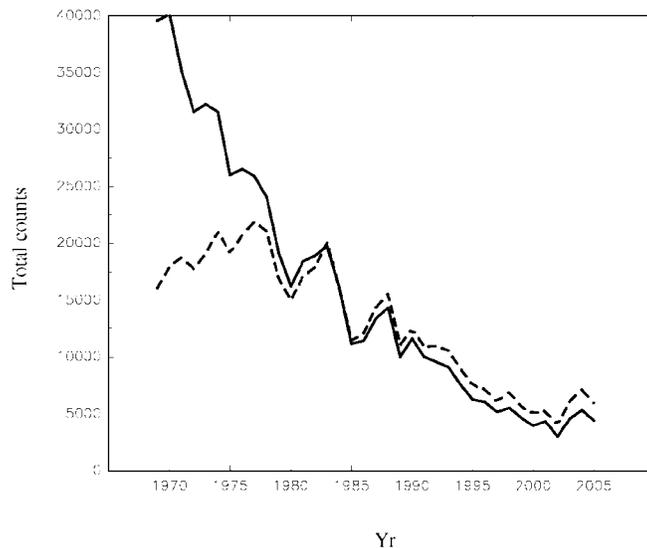


Figure 1. Total counts of northern bobwhite on Christmas Bird Count circles in United States and Canada, 1969–2005 (dashed line), compared to the total counts scaled to mean effort as indexed by number of party-hours (solid line).

The Noise term in equation 1 is normally distributed, with survey-specific variance and mean equal to zero. Given the efforts for standardization in the BBS, and the informal nature of the CBC, we anticipated that this error variance would be greater for the CBC than for the BBS. The inclusion of this random error term makes this an overdispersed Poisson regression because the variance of counts is greater than explained simply by the covariates. We implemented the analysis over many strata. In the model description below, we suppressed indexes for strata in the population effects, but we noted when random effects or hierarchical components of the model apply among strata.

Population effects.—We modeled the sampling effects for both surveys as departures from a zero baseline. Thus the remaining quantities in equation 1, the Population Effects,

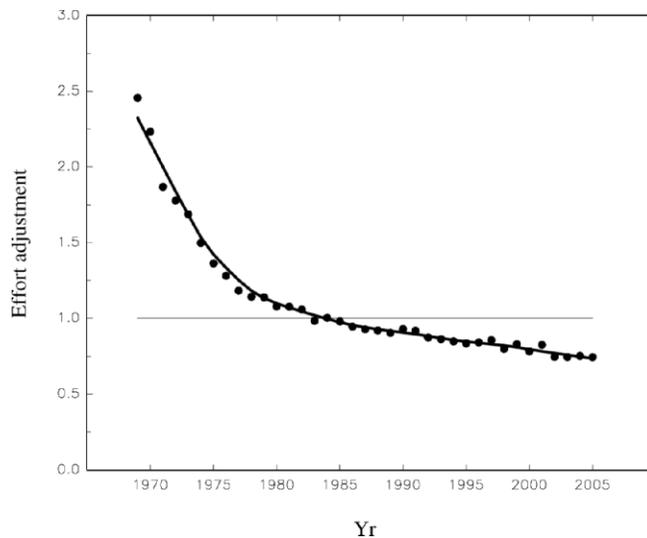


Figure 2. Simple effort adjustment for Christmas Bird Count circles in United States and Canada, 1969–2005.

can be thought of as a time series of numbers scaled to a typical log count for each survey. The Population Effects are free of sampling effects that may change through time, and that would be confounded with population change.

Let μ_y^k denote the population effects in year y for a count on survey k ($k = \text{BBS, CBC}$). Both sequences μ_y^{BBS} and μ_y^{CBC} ($y = 1, 2, \dots, Y$) reflect temporal population change, but $\mu_y^{\text{BBS}} \neq \mu_y^{\text{CBC}}$ for 2 reasons: 1) the 2 counts are conducted in different times of the year, when the population size is not the same; and 2) even if it were possible to conduct the surveys at the same time, the numbers of birds counted would be different due to basic issues of scale. In the bobwhite data, for instance, 23% of the BBS counts were zeros vs. 44% for the CBC. For nonzero counts, the mean and standard deviation for the BBS were 26.1 and 26.9, versus 34.1 and 49.0, respectively, for the CBC.

Nevertheless, the patterns of change in the 2 series are related. Let W_y denote the log of the proportional change in population size associated with the winter and spring season of calendar year y (roughly Jan to Jun), and let S_y denote the log of the proportional change in population size associated with the summer and fall season (roughly Jul to Dec). Recalling that the CBC is conducted in December and the BBS in June, we described temporal changes in counts over the surveys' annual cycles as

$$\mu_{y+1}^{\text{CBC}} = \mu_y^{\text{CBC}} + W_{y+1} + S_{y+1} \quad (2)$$

and

$$\mu_{y+1}^{\text{BBS}} = \mu_y^{\text{BBS}} + S_y + W_{y+1}. \quad (3)$$

Given that we can estimate μ_y^{BBS} and μ_y^{CBC} for $y = 1, 2, \dots, Y$, Eqs. 2 and 3 provide hope that we might be able to recover information about seasonal components of population change W_y and S_y .

We could not estimate the specific values of W_y and S_y because baseline differences in typical counts confound differences in survey methodology with differences in season of count. However, we could estimate

$$\delta_y^W = W_y - W_2$$

for $y = 3, 4, \dots, 36$, and

$$\delta_y^S = S_y - S_1$$

for $y = 2, 3, \dots, 36$ (details on why these parameter values are identifiable are presented in Link and Sauer 2007). Although the model does not allow estimation of the proportional change in the population for any particular winter or summer, it does allow comparisons among years. For instance, because we can estimate δ_5^W and δ_4^W , we can estimate $\delta_5^W - \delta_4^W = (W_5 - W_2) - (W_4 - W_2) = W_5 - W_4$, and compare the winter proportional change for years 4 and 5.

Hierarchical structure.—We used hierarchical structure in our model, that is, we treated parameters as random variables, themselves sampled from distributions governed

by higher-level parameters. The structure we used was the one implied by positing that, within strata, the effects W_y , $y = 1, 2, \dots, 36$, were sampled from normal distributions with stratum-specific variances $\sigma_s^2(W)$. Similarly, the structure was the one implied positing that the effects S_y , $y = 1, 2, \dots, 36$, were sampled from normal distributions with stratum-specific variances $\sigma_s^2(S)$. The quantities $\sigma_s^2(W)$ and $\sigma_s^2(S)$ can be estimated based on the variation in δ_y^W and δ_y^S .

We were interested in comparing the magnitude of $\sigma_s^2(W)$ and $\sigma_s^2(S)$; hence, we estimated

$$P_s = \frac{\sigma_s^2(W)}{\sigma_s^2(W) + \sigma_s^2(S)} \quad (4)$$

for each of the strata.

We extended the hierarchical structure one step further and supposed that $\sigma_s^2(W)$ and $\sigma_s^2(S)$ were sampled from lognormal distributions. This allowed us to estimate a composite version of P_s , namely

$$P = \frac{E(\sigma_s^2(W))}{E(\sigma_s^2(W)) + E(\sigma_s^2(S))}, \quad (5)$$

where $E(X)$ means the expected value of X .

Sampling Effects

Sampling effects are an example of what are commonly called nuisance parameters. These effects are of little interest per se but are required for valid inference about parameters of interest (in the present case, the population parameters μ_y^{BBS} and μ_y^{CBC} , and corresponding hierarchies of parameters). For bird surveys, the nuisance parameters often provide insights into the mechanics of counting birds, and have intrinsic interest for those designing surveys.

Sampling effects in the BBS.—We include 2 observer effects in our analysis of BBS observers. The first is a parameter η , included if the count was in the first year of the observers' service. Previous experience with BBS data suggests that η tends to be negative, with $\exp(\eta)$ often approximately 0.95, suggesting that new observers count 5% fewer birds than they would have under similar circumstances, but with more experience.

In addition, we include individual-specific observer effects, which we model hierarchically as having been sampled from a mean zero normal distribution with variance $\sigma^2(\text{obs})$.

Sampling effects in the CBC.—Many analyses of CBC data begin by scaling the counts by the corresponding effort (e.g., Butcher and McCulloch 1990). If we want to convert counts to a common scale corresponding to the mean level of effort, we might first divide them by the associated effort ξ , then multiply by the average effort $\bar{\xi}$. In terms of the loglinear model, this amounts to including a term of the form $\log(\xi/\bar{\xi})$ in equation 1.

Implicit in the usual scaling by effort is the notion that counts should increase at a constant rate as effort increases. This assumption is questionable because there is a fixed number of birds available for counting, so one might expect diminishing returns in numbers counted as effort increases.

Thus, Link and Sauer (1999) and Link et al. (2006) suggested a flexible family of alternative effort-effect models that includes $\log(\xi/\bar{\xi})$ as a special case; the general form of the effect of effort, on the log scale, is

$$f(\xi) = B\left(\left(\frac{\xi}{\bar{\xi}}\right)^p - 1\right)/p, \quad (6)$$

where B and p are parameters. Setting $B = 1$ and taking the limit as p approaches zero, one obtains the usual effort adjustment, $f(\xi) = \log(\xi/\bar{\xi})$. If $p < 0$ and B is positive, the effect of effort has an asymptote.

Johnson (1981) suggested scaling CBC counts by a factor of $1 - \alpha + \alpha \xi$, for some value α , $0 \leq \alpha \leq 1$. This represents a compromise between no effort adjustment (dividing by 1, if $\alpha = 0$) and scaling by effort (dividing by ξ , if $\alpha = 1$), and can be thought of as a linear approximation to the richer family of effort adjustments we used. This is because the linear Taylor approximation to $\exp(f(\xi))$ about $\xi = \bar{\xi}$ is $1 - B + B(\xi/\bar{\xi})$.

In our analyses of CBC data, we let the parameters p and B be stratum specific, modeled hierarchically as sampled from normal distributions.

Strata and Aggregation of Information

Historically, both surveys have been analyzed using physiographic strata, which reflect both methodological (e.g., variation in density of sample units) and bird-population (e.g., regional variation in bird abundance and change associated with regions) constraints on the analysis (Link and Sauer 2002, Link et al. 2006). We used Bird Conservation Regions within states and provinces as the fundamental scale for application of the model (e.g., Sauer et al. 2003, Link et al. 2006). We present within-stratum estimates of population change and annual abundances based on expected values of counts for randomly selected BBS routes (Link and Sauer 2002, Link et al. 2006). We also present area-weighted annual abundances for summaries among strata (Link and Sauer 2002).

We present results for 5 selected strata: Georgia Coastal Plain, Indiana Central Hardwoods, Maryland Mid-Atlantic Coast, Missouri Tall-grass Prairie, and Texas Oaks and Prairie. We provide composite results for the surveyed area. Presentation of BBS and CBC results on the same axis require that CBC results be scaled, and we scaled the CBC and BBS by adjusting the CBC counts to the mean level (on the log scale) of BBS counts for the region. We also analyzed survey results based only on BBS data for comparison with the combined survey results. We conducted BBS-only analyses using the methods described in Link and Sauer (2002). To assess the increase in efficiency associated with adding CBC data to the analysis, we compared the width of the credible intervals (Bayesian analog of confidence interval) from the analyses of the population (yr) effects, scaled to the median estimates.

Model Fitting

The model we described includes complex hierarchical structures and nonlinear parameters (parameters p in the

effort adjustment for the CBC), hence is not easily fit using conventional methods. However, it is easily fit using Bayesian methods. We conducted our analyses using program WINBugs, which is available for free download (<www.mrc-bsu.cam.ac.uk/bugs/>). The code we used to conduct our analyses is available from the authors on request.

Bayesian inference requires the specification of prior distributions for parameters; these probability distributions are a mathematical expression of existing knowledge about the parameters. Prior distributions can be informative (i.e., indicative that certain values of the parameters are more likely than others) or noninformative. On one hand, the existence of a formal mechanism for incorporating existing knowledge is an attractive feature of Bayesian inference. On the other hand, analysts often desire to let a single data set speak for itself and thus choose noninformative priors. Bayesian inference is based on posterior distributions, proportional to the product of the prior distribution and the likelihood. For a more detailed introduction to Bayesian inference, we suggest Gelman et al. (2004), Gilks et al. (1996), or Link et al. (2002).

We chose standard, noninformative priors on the highest-level parameters. By highest-level parameters we mean 1) those governing the distributions of hierarchically modeled parameters, such as the mean and variance of the distribution of effort parameter p across strata; and 2) parameters defined without hierarchical structure (e.g., the first-yr observer effect η). These priors were either flat normal distributions (for means) or flat gamma distributions (for variances).

Program WINBugs implements Bayesian inference using Markov chain Monte Carlo (MCMC; Spiegelhalter et al. 1999). Markov chain Monte Carlo evaluates features of posterior distributions through simulation. Prior distributions and data are used to generate series of correlated samples (Markov chains, or chains) from the posterior distribution. We used posterior medians as point estimates, and the 2.5th and 97.5th percentiles to define interval estimates.

Care must be taken in implementing MCMC. First, it must be recognized that early values of the chains may not be representative samples of the posterior distribution. This problem is avoided by discarding an appropriately large number of early observations, described as a burn-in. A second consideration is that the reliability of the simulations depends on the degree of autocorrelation in the chains. Analysts need to take care that the combination of chain length and autocorrelation are sufficient for reliable inference.

We generated chains of length 105,000, discarding the first 5,000 values as a burn-in. The autocorrelation of the resulting chains was limited relative to the length of the chains, indicating satisfactory performance of the Markov chain simulation. Owing to the complexity of the model and the large size of the data set (32,281 counts by 5,353 observers for the BBS, and 23,417 counts on 1,134 circles

Table 1. Regional and composite trend estimates (% change/yr) with associated 95% credible intervals, and estimated relative magnitude of seasonal components of population variation (P_s) and its standard error.

Region	Trend	95% credible interval	P_s	SE
Georgia Coastal Plain	-4.71	-5.38, -4.06	0.56	0.30
Indiana Central Hardwoods	-4.62	-5.75, -3.46	0.91	0.19
Mid-Atlantic Coast	-7.62	-8.26, -6.98	0.21	0.27
Missouri Tall-grass Prairie	-2.53	-3.39, -1.65	0.86	0.24
Texas Oaks and Prairie	-5.63	-6.30, -4.97	0.45	0.25
Survey-wide	-3.56	-3.80, -3.32	0.75	0.16

from the CBC), the simulation took approximately 10.5 hours on a 3.2-GHz desktop computer system.

RESULTS

Nuisance Parameters

As we anticipated, the CBC data were noisier than the BBS, in the sense that the standard deviation of the noise term in equation 1 was larger for the CBC data than for the BBS data. The posterior median (a Bayesian point estimate) of the CBC error standard deviation was 1.459, with 95% credible interval (1.435, 1.482). For the BBS, the estimate was 0.322, with credible interval (0.317, 0.328). The standard deviation of observer effects for the BBS was 1.067, with credible interval (1.041, 1.093), and the standard deviation among CBC circle effects was 1.494, with credible interval (1.418, 1.573).

A small but significant portion of the variability among BBS observers can be attributed to change in the pool of observers: we calculated the posterior distribution of the slope of the least squares line regressing observer effects on the observer's first year of service, obtaining a point estimate of 0.012 (0.008, 0.016). This small amount of trend in the observer pool, combined with the pattern of change among observers in the data set, is sufficient to account for a positive bias of approximately 0.9% per year in estimating population change, underscoring the importance of controls for difference among observers.

We estimated the first-year observer effect for the BBS as -0.065 , with interval estimate $(-0.084, -0.047)$. This means that observers typically counted approximately $6.3\% = 100\% (1 - \exp[-0.065])$ fewer bobwhites than would otherwise be expected in their first year of service. Failure to include this effect could produce a positive bias in trend estimate.

The effort effects modeled for the CBC ($f(\xi)$, defined by eq 6) had mean value of p estimated as -0.168 $(-0.319, -0.004)$ and mean value of B estimated as 0.827 $(0.706, 0.954)$. The pair $(p = 0, B = 1)$ corresponding to the usual effort adjustment (i.e., scaling count by effort) was in the extreme tail of the joint posterior distribution, with $<0.83\%$ of the distribution further away from its center. The value 0.83% is a Bayesian p -value strongly indicating the inadequacy of the usual effort adjustment.

Population Effects

Sample results from the 5 strata and the summary of 69 strata all show declining populations (Fig. 3). All strata

exhibit declines, but many also show large declines in counts in the mid-1970s, a period of notably severe winters in the eastern and central United States (Link and Sauer 2007).

Comparison with a BBS-only analysis indicates only a slight increase in precision associated with use of the composite analysis. Median half-width of the confidence interval in the composite summary was 6.43% of the point estimate; in an analysis of the BBS data alone, the typical 95% credible interval was estimated $\pm 7.50\%$ of the point estimate.

We defined population trend as the geometric mean percentage population change per year over the time interval; estimates for 5 strata are given in Table 1. The 5 regions we summarized and the survey-wide average document significant declines. The area- and abundance-weighted average across the range of the species was -3.56% per year $(-3.80, -3.32)$.

Estimates of the relative magnitude of winter variation P_s defined by equation 4 are also given in Table 1. The composite value P defined by equation 5 was more somewhat more precisely estimated as 0.747 (0.162); 90.1% of the posterior mass for this parameter was above 0.500, indicating reasonable support for the conclusion that northern bobwhite populations tend to have larger fluctuations in the winter and spring portion of the year than in the summer and fall portion of the year.

DISCUSSION

Although many analyses have documented declines in bobwhite populations (e.g., Sauer et al. 2005a), integrating the primary monitoring databases for the species provides an opportunity for us to enhance the quality of the estimates of population change and evaluate seasonal patterns of change. Our composite analysis of BBS and CBC data provides composite results of slightly higher precision, although the large variation associated with CBC data tends to mitigate their value in improving the precision of the overall estimates of change.

The BBS and CBC data sets do not support estimation of actual population size but can be used to estimate population change over their staggered annual cycles. Similarly, composite analysis of the BBS and CBC does not support estimation of baseline differences in the proportions counted by the 2 surveys, these being confounded with temporal change in the population between the summer and winter. Consequently, it is not possible to estimate year-specific seasonal rates of population change. However, under the

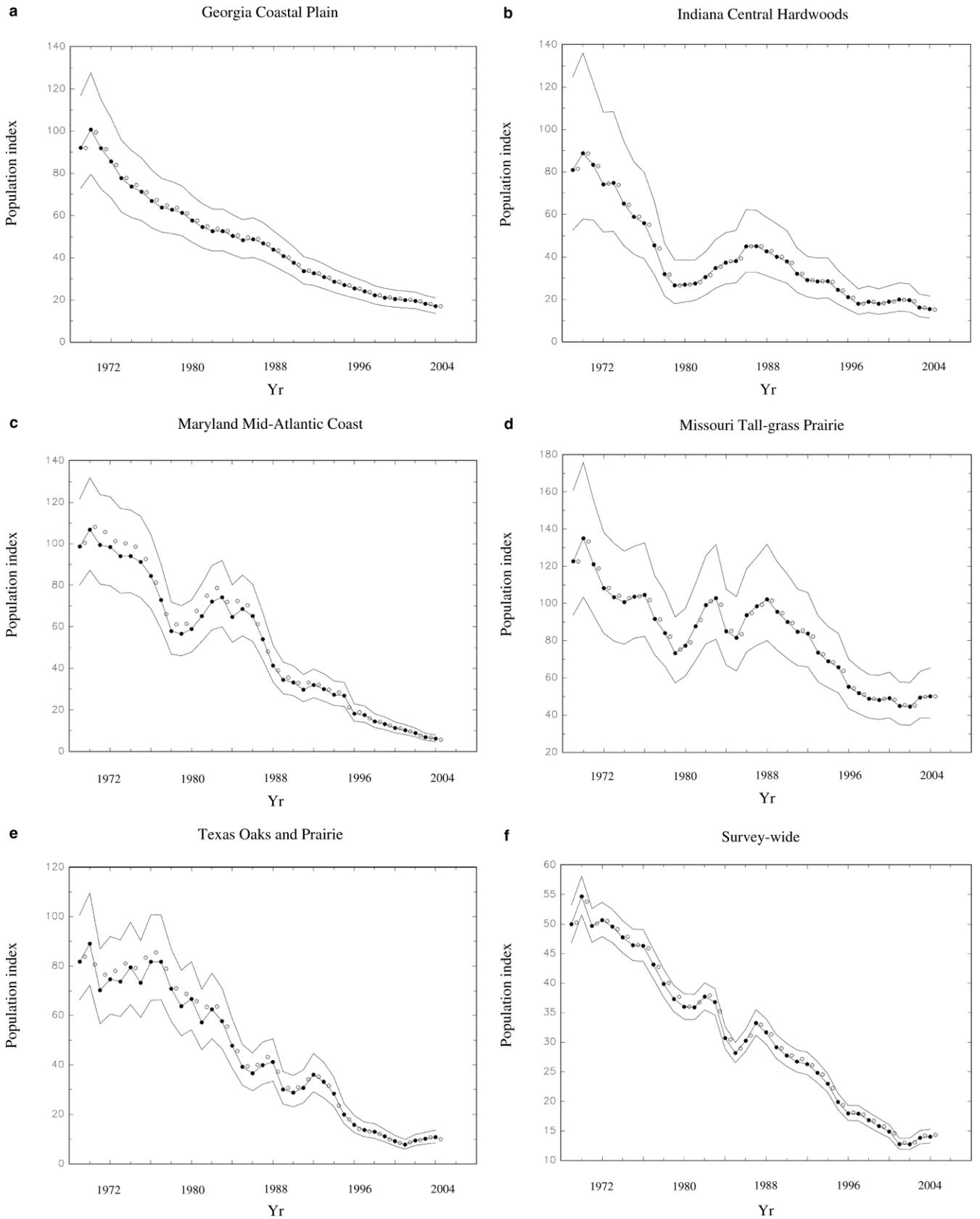


Figure 3. Expected counts of northern bobwhite on North American Breeding Bird Survey (BBS) routes, 1972–2004, in (a) Georgia Coastal Plain, (b) Indiana Central Hardwoods, (c) Maryland Mid-Atlantic Coast, (d) Missouri Tall-grass Prairie, (e) Texas Oaks and Prairies, and (f) survey-wide, all adjusted for observer effects (solid dots and line), with 95% credible intervals. Open circles are expected counts from Christmas Bird Count circles, adjusted for effort and scaled to mean level of BBS counts.

model we presented, we can estimate variation in seasonal rates of population change.

A critical result from the composite analysis is that most of the variation in population size is associated with the winter to spring period. Evaluation of population dynamics by intervals within the year has great potential as supplemental information for validating or directly tying the monitoring information to demographic parameters and models (e.g., Guthery et al. 2000). Population change occurring in winter to spring reflects changes in survival, whereas change from summer to fall reflects both reproduction and survival. Unfortunately, the timing of the BBS and CBC makes it difficult to directly connect the survey data to major events influencing survival and reproduction. Bobwhites are already reproducing when the BBS is conducted in June (although young are not counted in the survey), and the CBC counts occur during the middle of most hunting seasons, which span from October to March. Consequently, timing of demographic modeling exercises may not coincide with the survey times. However, the CBC counts generally occur before severe winter events; hence, the variation associated with the winter to spring portion of the year likely reflects the winter mortality in bobwhite associated with severe winters. Declines in the severe winters of the mid-1970s are clearly seen in the population time series for the northern strata (Fig. 3). Although it is beyond the scope of our study, we note that the seasonal components of population we presented can be used to evaluate hypotheses regarding the factors influencing population change.

Analysis of the nuisance factors provides many insights into the mechanics of surveys, showing that both effort in counting (for the CBC) and observer startup and temporal variation play a large role in influencing our views of population change. Changes in levels of nuisance factors over time increase the counts. Startup effects cause counts to be 6.3% lower the first year of counting, and changes in observer quality introduce a positive bias of about 0.9% per year in estimates of change. Effort effects on CBC counts are well known (Butcher and McCulloch 1990), but our analysis shows that standard effort adjustments (e.g., simple division by effort as implemented in many analyses) are not sufficient for analysis of bobwhite data.

Other counting factors may also influence estimates of population change in bobwhites. The tendency for variation in calling rates with abundances and other environmental factors (Wellendorf et al. 2004) has long been thought to cause underestimates of population change. The hierarchical model we described could be modified to include a model for variation in calling rates if sufficient information were available for the regions used in the analysis.

MANAGEMENT IMPLICATIONS

Composite analyses of BBS and CBC data permit efficient use of information collected in 2 seasons for estimation of population change. The composite model also provides a general framework for associating factors that may influence bobwhite populations, such as habitat changes and harvest,

with seasonal population change and baseline abundance. We encourage the use of this model in developing the predictive models needed to assess the consequences of possible management actions on bobwhite populations, and suggest that the composite survey results could be used directly as a state variable for assessing the consequences of changes in harvest rates and other management actions.

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