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Abstract.

Using an improved testing procedure based on bootstrap and weighted jackknife confidence intervals with the same model as used in Bechtold et al. (1991) and Ruark et al. (1991), analysis in this paper generally confirm the results of a significant decrease in growth rate in pine in Georgia and Alabama for 1972 - 1982 (5th cycle) relative to 1961 - 1972 (4th cycle) discussed in these papers.

KEY WORDS. Pine growth, decline, bootstrapping, weighted jackknife, FIA

Introduction

Since 1928, the Forest Inventory and Analysis units (FIA) of the USDA, Forest Service, have inventoried the forest resources of the U.S. These surveys have generated estimates of aggregates, such as area in major land classes and/or forestry type volumes by tree species, and changes in areas and volumes over time.

FIA data are partially designed to assess change in the forest resource over large areas. FIA inventories in recent decades have revealed a decrease in the total timber resource of southeast U.S.A., a finding supported by the work of Sheffield et al. (1985) and Sheffield and Cost (1987). This decline is as much as 17-23% for natural loblolly pine and 27% for natural shortleaf pine (Zahner et al. 1989, Bechtold et al. 1991). Knight (1987) attributed the decline to four factors -- declining area of timberland, inadequate regeneration after harvest on nonindustrial private forest lands (NIPF), increased tree mortality, and reduction in the rate of tree and stand growth.

Ruark et al. (1991) used FIA data to compare the periodic annual increment in basal area of selected naturally regenerated pine stands throughout Alabama and Georgia. Estimated growth rates between 1972 and 1982 (5th cycle) were compared with estimated growth rates obtained during the previous 10-year survey cycle (4th cycle). Separate analyses were conducted for loblolly, longleaf, shortleaf, and slash pine cover types. Comparisons of growth rates yielded reductions ranging from 3 - 31% in both states. All results were statistically significant except for the 3% decline in natural loblolly pine in Alabama. The agent(s) causing the decline were not identified.

The purpose of this paper is to focus on improving the tests of significance used in Bechtold et al. (1991) and Ruark et al. (1991) for assessing the significance of the growth decline.

Basic Approach to the Problem

A brief review of the history of the growth rate decline issue in the southern states seems relevant. Sheffield et al. (1985) first reported a reduction in growth rate of yellow pines. Subsequently the possible effect of changing stand dynamics on growth rate decline was examined (Bechtold et al. 1991). This argument led to the introduction of a model in the study to explore the relation between the growth rate and stand dynamics. Bechtold, et al. (1991) and Ruark et al. (1991) developed proper models based on the data collected by the U.S. Southeastern and Southern FIA units. Bechtold et al. (1991) used t-tests to test the significance of the so called adjusted growth rates which are the predicted growth rates based on the model, while Ruark et al. (1991) used F-tests to test if the "population marginal means" or "least squares means" in these two survey cycles were equal or not.

Both of these two approaches can be improved upon. The main drawbacks of the Bechtold et al. (1991) approach are:

- They used t-tests without checking the independence of the adjusted growth rates which are certainly not independent;
- The adjusted growth rates used did not completely eliminate the effect of stand dynamics, since the adjusted growth rates were obtained as predicted growth rates from the regression equations. Hence, these growth rates still reflect differences in stand conditions between cycles.

Ruark et al. (1991) was aware of these drawbacks and used F-tests to test the hypothesis that the "population marginal means" are equal, based on their model. But they obtained the population marginal means by replacing all covariates by their means in the fitted regression equation (SAS Manual, SAS/STAT guide for personal computers, Version 6 Edition), hence, they also failed to eliminate the stand dynamics completely.

A better approach is as follows: Consider the model given by Bechtold et al. (1991):

$$ln(G) = b_0*c_1+b_1*c_2+b_2*S+b_3*\ln(A)+b_4*l_n(N)+b_5*P+b_6*\ln(M+1)+\epsilon_1 \qquad (1)$$

where G is the gross growth rate of Pines (denoted by "Pine") or all species (i.e., pine and non-pine combined, denoted by "All") with diameters larger than or equal to 1"; c1, c2 are indicator variables of the survey cycles 1961-1972 and 1972-1981 respectively; A is the age of the stands; N is number of pines with 1" dbh (diameter at breast high) and larger; S is site class; P is the ratio of pine basal area to total basal area; M is the pine basal area mortality; and E is an error term with mean zero and variance σ^2. The basic assumption underlying this model is that effects of the covariates on the growth rate do not change with cycle, so that any difference in the growth rate not due to these covariates will be picked up by b_0 and b_1. To test the hypothesis that gross growth rates of these two cycles are equal, with the effects of all other independent variables being eliminated, we should simply test

$$\text{Ho:} \quad b_0 = b_1.$$

If $b_0 \neq b_1$, the growth rates between these two cycles are different, even though the stand dynamics are the same; but if $b_0 = b_1$ the gross growth rates will be equal if the stand dynamics are the same.

The Working Models

It is helpful to briefly introduce the models developed by Bechtold et al. (1991) for the Georgia analysis. In this study, the pines were classified into three classes; loblolly, shortleaf pine and slash pine. For each kind of pine, two cases were considered separately: only the pine component, "Pine" and all trees, "All", e.g., for a loblolly pine forest, considering loblolly pine only and considering all trees. Thus, six cases were analyzed based on model (1).

Since data in Georgia and Alabama are collected in somewhat different ways and with somewhat different variables being measured, a slightly different model had to be used for the joint analysis. The model used in this analysis was

$$\ln G(i)=b_0c_1+b_1c_2+b_2\ln(QMD)+b_3 lnN+b_4S+b_5P+b_6\ln(M+1) \quad , i=1,2 \tag{2}$$

where G(1) = basal area growth of pine trees 1" and larger, G(2) = basal area growth of pine trees 5" and larger, QMD = initial quadratic mean diameter of pine trees 1" and larger, and other variables are as defined for model (1). The model was fit to data for loblolly pine, longleaf pine, shortleaf pine, and slash pine (Georgia only).

For both data sets only a few plots were remeasured in both the 4th and 5th cycle so that correlation of growth measurements over time should be fairly insignificant. For the current analysis we used models (1) and (2) as used by Bechtold et al. (1991) and Ruark et al. (1991).

Besides the five independent variables considered in model (1), other variables such as pollution, weather and rainfall may be involved in the growth rate of pines. Even though these variables and some other variables might be very important to the growth rate, such data are not collected by FIA units. The effect of omitting such variables is that the model error will usually be increased. From the view point of application, this model error is fairly measured by R^2.

Statistical Analyses

To conduct statistical tests of the hypothesis

$$H_0: b_0=b_1;$$

one approach is to use the classical linear model theory to perform the test on H_0, assuming model (1) is reasonable. But, some residual plots suggest that the model might have heterogeneous error, and the normal probability plots are not satisfactory either. This suggests that the F-test used in the classical linear model theory may not be appropriate so more robust procedures need to be considered.

Two such procedures are bootstrapping and jackknifing. These two methods are nonparametric and generally applicable. To use bootstrapping and jackknifing to test the hypothesis H_0: $b_0=b_1$, we need to construct confidence intervals, then see if the confidence intervals contain zero to accept H_0, or reject H_0 if zero is not contained in the interval.

Bootstrapping and jackknifing have another advantage. We know the data set was first used to select a proper model, and then used to fit the model. It is well known this might cause over-fit to the model (see, for example, Efron (1982)). The overfit to the model means the sum of squares of the residuals is too small compared to the model error variance. By using bootstrapping and jackknifing (delete say, 5 or 15 units each time), a different data set is obtained for each iteration, resulting in a better test using the selected model. This idea is similar to the idea of using bootstrapping and jackknifing in cross-validation.

Description of the Bootstrap and Jackknife Methods

To construct bootstrap confidence intervals for b_0-b_1, we:

- Draw a bootstrap sample of size n from the observation $\{(y_i, x_{1i}, \ldots, x_{ki}), i = 1, \ldots, n$ and k = # of covariates$\}$ by using simple random sampling with replacement;

- Use the bootstrap sample to fit the model (1), and then obtain $\hat{b}_0-\hat{b}_1$;

- Repeat the first two steps B times to obtain an empirical distribution of $\hat{b}_0-\hat{b}_1$;

- Find the α and $1-\alpha$ percentiles from the empirical distribution, say $CDFB^{-1}(\alpha)$ and $CDFB^{-1}(1-\alpha)$. Then the $1-2\alpha$ central confidence interval for $b_0 - b_1$ is $[CDFB^{-1}(\alpha)]$, $CDFB^{-1}(1-\alpha)]$.

If $0 \notin [CDFB^{-1}(\alpha), CDFB^{-1}(1-\alpha)]$ then the hypothesis H_0 is rejected with significance level 2α.

This bootstrap confidence interval is recommended by Efron and Tibshirani (1986) as a robust procedure used for regression models.

A more robust procedure which was specially designed for heterogeneous error regression models but without "neglecting the unbalanced nature of the regression data $\{y_i, \underline{x}_i\}$", was proposed by Wu (1986). We also use this weighted jackknife procedure as follows:

Suppose model (1) is rewritten in matrix notation

$$y=X\beta+e \tag{3}$$

where $y = (y_1, ..., y_n)'$ is the observations on ln(G); $X = [X_1, ..., X_n]'$ are the observations on (C_1, C_2, s, ln(A), ln(N), P, ln(M+1); $e = (e_1, ..., e_n)'$ is the error term; $\beta = (b_0, b_1, b_2, ..., b_6)'$ is a (7x1) vector of parameters; and e_i and e_j are uncorrelated with mean 0, for $i \neq j$. Let $\hat{\beta}$ be the LSE of β based on the data. For a subset $s \subset \{1, ..., n\}$; let $y_s = (y_i, i \epsilon s)'$; where y_s is arranged in the same order as y. Similarly, we define X_s and e_s. From the subset s, we have the following s-model

$$y_s = x_s \beta + e_s. \tag{4}$$

Let $\hat{\beta}_s$ be the LSE of β based on the s-model. From $\hat{\beta}_s$ and $\hat{\beta}$, we can obtain the corresponding $b_{0s} - b_{1s}$ and $b_0 - b_1$. For a fixed sample of size $r \geq 7$, the jackknife confidence intervals can be constructed in the following way:

- Draw a jackknife sample s of size r randomly without replacement
(Since the number of all possible samples deleting n-r observations at a time is quite large for most r values, J such samples are generated, deleting n-r units at random. We used J = 1000.);

- Use the jackknife sample s to fit model (3), and then obtain
$\tilde{b}_{0s}-\tilde{b}_{1s}=(\hat{b}_0-\hat{b}_1)+\frac{(r-k+1)^{1/2}}{(n-r)}[(\hat{b}_{0s}-\hat{b}_{1s})-(\hat{b}_0-\hat{b}_1)]$;

- Repeat the first two steps J times, then J observations $\tilde{b}_{0sj} - \tilde{b}_{1sj}$ (j = 1, ..., J) are obtained. Assign weight $\left|X_s X_{si} / \sum_{j=1}^{J} \left|X_{sj} X_{sj}\right|$ to observation $\tilde{b}_{0s}-\tilde{b}_{1s}$ and then obtain an empirical distribution of $\tilde{b}_{0s}-\tilde{b}_{1s}$ based on the given weight;

- Find the α and $1-\alpha$ percentiles from the empirical distribution, say $CDFJ^{-1}(\alpha)$ and $CDFJ^{-1}(1-\alpha)$. Then the $1-2\alpha$ central confidence interval for $b_0 - b_1$ is $\{CDFJ^{-1}(\alpha), CDFJ^{-1}(1-\alpha)]$

The test of H_0 based on the jackknife confidence interval is similar to the one based on the bootstrap confidence interval.

Results and Conclusions

Three approaches to analyze the Georgia and Georgia-Alabama FIA data have been described. For the classical linear model theory approach to perform the test on H_0 we use the GLM procedure in SAS. For the bootstrap approach and the jackknife approach, we wrote a Fortran program supported by an IMSL subroutine to form confidence intervals for $b_0 - b_1$. The R^2 of the models and the estimates of b_0 and b_1 are given in Table 1.

Table 1. R^2 and estimates of b_0 and b_1

	Loblolly		Shortleaf		Slash	
	Pine only	All species	Pine only	All species	Pine only	All species
R^2	.542	.396	.598	.440	.389	.326
$\hat{b}_0$	-1.258	-.610	-1.815	- .811	-1.356	- .794
$\hat{b}_1$	-1.468	-.785	-2.704	-1.661	-1.680	-1.066

From Table 1, we can see that $\hat{b}_0 > \hat{b}_1$ in all cases. To test the hypothesis $H_0: b_0 = b_1$ for the Georgia data, Table 2 shows the confidence intervals formed by bootstrapping and weighted jackknifing with five plots deleted at a time with significant levels .01 and .05 for 1000 samples and the observed significance levels (p-values) for the classical test. Table 3 shows the corresponding values for the Georgia-Alabama data for basal area of pine trees 5" and larger (PSG5) and pine trees 1" and larger (PSG1). It may be pertinent to point out that the R^2 values given in Table 1 are low which is typical in analyzing FIA data. Low R^2 result in big sums of squares of residual which tends to reduce the F value of a test. As we will see in the following, most of the test results are significant. Thus, our analyses are less affected by the low R^2 values.

Table 2. Confidence intervals of $b_0 = b_1$ using weighted jackknifing deleting 5 plots at a time and bootstrapping using 1000 samples and classical p-values, Georgia 4th and 5th cycle

Tree Type	C.I.	Plot Type	Bootstrap	Jackknife-5	p-values of classical analysis
Loblolly Pine	95%	All	(.073, .279)	(.074, .272)	0.0002
	99%	All	(.035, .296)	(.029, .306)	
	95%	Pine	(.105, .320)	(.110, .317)	0.0012
	99%	Pine	(.073, .349)	(.068, .360)	
Slash Pine	95%	All	(.079, .459)	(.090, .464)	0.0001
	99%	All	(.018, .501)	(.018, .514)	
	95%	Pine	(.107, .528)	(.105, .514)	0.0001
	99%	Pine	(.044, .592)	(.033, .581)	
Shortleaf Pine	95%	All	(.115, .459)	(.076, .473)	0.0038
	99%	All	(.068, .516)	(.023, .534)	
	95%	Pine	(.084, .467)	(.056, .484)	0.0020
	99%	Pine	(-.013, .512)	(-.026, .557)	

Table 3. Confidence intervals of $b_0 = b_1$ using weighted jackknifing deleting 5 plots at a time and bootstrapping using 1000 samples and classical p-values, Georgia-Alabama 4th and 5th cycle

State & Tree Type	C.I.	Model	Bootstrap	Jackknife-5	p-value of classical analysis
Alabama Loblolly	95%	PSG1	(.025, .232)	(.019, .234)	0.0182
	99%	PSG1	(-.014, .264)	(.001, .277)	
	95%	PSG5	(-.008, .172)	(-.063, .112)	0.5754
	99%	PSG5	(-.045, .204)	(-.091, .143)	
Alabama Longleaf	95%	PSG1	(.172, .503)	(.180, .486)	0.0001
	99%	PSG1	(.119, .550)	(.129, .523)	
	95%	PSG5	(.084, .363)	(.055, .333)	0.0061
	99%	PSG5	(.038, .412)	(.011, .361)	
Alabama Shortleaf	95%	PSG1	(.044, .321)	(.031, .312)	0.0179
	99%	PSG1	(-.009, .362)	(-.014, .368)	
	95%	PSG5	(.106, .432)	(-.009, .277)	0.0742
	99%	PSG5	(.061, .462)	(-.072, .366)	
Georgia Loblolly	95%	PSG1	(.235, .602)	(.170, .337)	0.0001
	99%	PSG1	(.159, .642)	(.124, .366)	
	95%	PSG5	(.158, .324)	(.121, .282)	0.0001
	99%	PSG5	(.131, .351)	(.086, .338)	
Georgia Longleaf	95%	PSG1	(.014, .275)	(.016, .267)	0.0001
	99%	PSG1	(-.027, .306)	(-.033, .304)	
	95%	PSG5	(-.030, .227)	(-.002, .244)	0.0512
	99%	PSG5	(-.091, .253)	(-.049, .281)	
Georgia Shortleaf	95%	PSG1	(.275, .517)	(.257, .514)	0.0001
	99%	PSG1	(.223, .552)	(.219, .549)	
	95%	PSG5	(.214, .523)	(.172, .433)	0.0001
	99%	PSG5	(.184, .563)	(.116, .472)	
Georgia Slash	95%	PSG1	(.125, .379)	(.122, .396)	0.0001
	99%	PSG1	(.079, .409)	(.062, .453)	
	95%	PSG5	(.053, .310)	(.032, .296)	0.0063
	99%	PSG5	(.020, .354)	(-.017, .357)	

For the Georgia data, the bootstrap test and the weighted jackknife test showed significant growth differences at both the $\alpha = .05$ and $\alpha = .01$ levels in 11 out of 12 cases. The classical result showed significant differences in all cases. For the Georgia-Alabama data, the classical test, the bootstrap, and the weighted jackknife generally found significant differences for species between cycles. The bootstrap only found non-significant differences for loblolly pine - Alabama at the $\alpha = .01$ level for stems > 1", and at both the $\alpha = .01$ and .05 level for stems > 5"; for shortleaf pine in Alabama at the $\alpha = .01$ level for stems > 1", and at both levels for stems > 5". The weighted jackknife found a few more nonsignificant differences. The same ones were significant as with bootstrapping with the exception that the weighted jackknife also found a significant difference at the $\alpha = .05$ level for Georgia longleaf pine for stems > 5". In addition, weighted jackknifing found nonsignificant differences for shortleaf pine in Alabama at both levels for stems > 5" and for slash pine in Georgia for stems > 5" at the $\alpha = .01$ level.

The results of our analyses confirmed the growth decline in pine in Georgia and Alabama. The use of the robust bootstrap and jackknife procedures made our results more reliable than similar results in earlier papers.

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