Tag-Based Estimates of Annual Fishing Mortality of a Mixed Atlantic Coastal Stock of Striped Bass

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Abstract.—Tag-based estimates of annual survival and fishing mortality rates supplement annual stock assessments of migratory striped bass *Morone saxatilis* in the interjurisdictional fishery along the Atlantic coast. We estimated a 17-year time series of annual survival and fishing mortality ($F$) rates for striped bass (>$711$ mm) tagged during winter trawl studies (1988–2004) off the coasts of North Carolina and Virginia. The geographic and temporal distributions of tag recoveries were consistent with published patterns of striped bass migration and indicated that this southern overwintering aggregate of striped bass is composed of mixed stocks. Incremental increases in bias-adjusted annual fishing mortality rates (from $0.00$–$0.26$) and decreases in the proportion of fish released alive (from $0.762$–$0.198$) coincided with periods of regulatory change during the 17-year time frame. Our estimates of $F$ fall below the current management triggers and should be considered along with other estimates of $F$ within the striped bass management process.


Managers set and change harvest regulations for the Atlantic striped bass fishery based in part on estimates of annual fishing mortality rates ($F$). The current FMP (Amendment 6) stipulates threshold ($F = 0.41$) and target ($F = 0.30$) rates of fishing mortality (Beal et al. 2003). The current threshold $F$ was derived from an estimate of the maximum sustainable yield, and the target rate was designed to conservatively reduce adverse effects on spawning potential (Beal et al. 2003). Estimates of $F$ are obtained through virtual population analysis (Shepherd 2002) and tag-based analysis (Smith et al. 2000).

Initially, coastal striped bass tagging programs evaluated efforts to restore Atlantic stocks of striped bass (Wooley et al. 1990), but currently tag recoveries, including estimates of annual fishing mortality rates, are used for striped bass management. The U.S. Fish and Wildlife Service (USFWS) manages the coastwide tag-

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recovery database, and a reward system (cash or baseball caps) and toll-free phone number encourage tag reporting by commercial and recreational fishermen. Tag number, date and location of capture, and disposition of fish (killed or released), which are needed for estimating annual survival and fishing mortality rates, are requested during toll-free interviews with tag-reporting commercial or recreational fishermen.

Eight tagging programs overseen by the ASFMC Striped Bass Tagging Subcommittee (SBTS) provide estimates of $F$ for striped bass stock assessments by spawning stock (Rappahannock River, Virginia; Chesapeake Bay, Maryland; Delaware River, and Hudson River) and mixed coastal stock (North Carolina and Virginia; New Jersey; Long Island, New York; and Massachusetts). The SBTS assumes that striped bass greater than 711 mm (TL) at time of tagging are fully recruited to the coastal migratory population and estimates $F$ and annual rates of survival ($S$) for this size-group from the tagging data (Smith et al. 2000). Our objectives are to describe the current ASFMC striped bass tagging analysis protocol and estimates of $F$ and $S$ within a 17-year time series (1988–2004) of striped bass tagged off the coasts of North Carolina and Virginia during the cooperative winter tagging cruises.

### Methods

**Cooperative Winter Tagging Cruise**

During winter periods (mid-January to mid-February) of each year from 1988 to 2004, a total of 8,748 striped bass 711 mm TL and larger were captured with trawls and tagged on overwintering grounds off North Carolina and Virginia (Figure 1; Table A.1 in the appendix). Striped bass were tagged with sequentially numbered internal anchor tags (inserted through a small incision in the abdominal cavity) and released near the capture location. Numbers of annual tag releases and tag recoveries of these striped bass were queried from the USFWS database and formatted as a triangular matrix (Table A.1).

**Analysis Methods and Models**

**Distribution of harvest.**—The geographic and temporal distributions of tag recovery data were examined by descriptive methods (tabulation by month of recovery and state of recovery). With a descriptive approach, one can follow temporal changes in location of tag recoveries and assess whether migratory patterns of the tagged cohort are consistent with those published for untagged fish. Further, one can see whether the tagged cohort represents a single or mixed stock based on recoveries on spawning grounds during spring.

We used a chi-square test of nonmixing (as suggested by Latour et al. 2001) to examine mixing among newly and previously tagged striped bass within three geographic areas of tag recovery and four recent periods. The three geographic areas of tag recoveries were Chesapeake Bay, waters off New Jersey and southward excluding Chesapeake Bay, and waters off New York and northward. Overlap of the four periods (1999–2001, 2000–2002, 2001–2003, and 2002–2004) allowed for complete assessment of nonmixing from 1999 to 2004, and the short 3-year periods reduced the numbers of observations per cell. The observations of geographic and temporal tag recoveries and the chi-square tests of nonmixing were in support of analyses of tag-based estimates of $F$ (a complete analysis of tag-based migration rates and migratory behaviors was outside of our study objectives).

The striped bass tag recovery data were analyzed in program MARK (White and Burnham 1999), where survival rates were derived from a suite of Seber (1970) models and assumptions followed Brownie et al. (1985). Model selection followed an information-theoretic approach, based on Kullback–Leibler information theory and Akaike’s information criterion (AIC; Burnham and Anderson 2002). Twelve candidate models represented alternative hypotheses of temporal change in annual survival ($S$) and recovery probability ($r$) during the 17-year time frame, where $r$ is the probability that tags from dead, marked individuals are reported during each period between releases, regardless of the source of mortality (as distinguished from the reporting rate, $\lambda$, of anglers and commercial watermen; see the section on bias-adjusted estimates
of survival below). Candidate models categorize four groups: time-specific, constant, monotonic trend (with time as a covariate), or regulatory period-specific (Table 1). The global model, a time saturated model, parameterized year-specific survival rates, and the two least parameterized models represented survival as constant during the time series. Four regulatory period models parameterized survival as constant within time intervals between regulatory changes. Regulatory periods were defined based on the following four previously described periods of management changes: the restricted harvest period (1988–1989), the interim fishery (1990–1994), and two full-fishery periods (1995–1999, and 2000–2004). Given the importance of recent years (2003 and 2004) within the terminal regulatory period (2000–2004), we also modeled survival of the terminal year separately and of the most recent 2 years separately. Three models parameterized time as a covariate and reflect hypotheses of increasing or decreasing monotonic trends in survival.

**Diagonal procedures.**—Model adequacy is a major concern when deriving inference from a model or a suite of models. Overdispersion, inadequate data (such as low sample size), or poor model structure may cause a lack of model fit. In striped bass tagging data, lack of independence among recaptures can result from schooling behavior and causes overdispersion (i.e., observed data are more dispersed than expected under the global model). Overdispersion causes variance to be underestimated and models to be overparameterized. An estimate of the variance inflation factor (i.e., \( \hat{c} \)) adjusts for overdispersion (after adjustment, the corrected AIC \( [AIC_c] \) is called QAIC\(_c\); Anderson et al. 1994). Overdispersion and model fit statistics were estimated with the global model. We estimated \( \hat{c} \) by dividing the observed Pearson chi-square value (goodness-of-fit statistic of the global model) by the expected Pearson chi-square value (derived via MARK from a 1,000-replicate bootstrap analysis of the global model). The goodness-of-fit probability of the global model was examined with a bootstrap-derived P-value based on model deviance (Burnham and Anderson 2002). A low P-value and a large estimate of \( \hat{c} \) (>4), in part, imply inadequate model structure (Burnham and Anderson 2002). A low bootstrap-derived P-value combined with a moderate estimate of \( \hat{c} \) (>1 but <4) supports overdispersion rather than an inadequate model structure. Overdispersion is correctable with \( \hat{c} \) adjustment (as described above). Correcting for inadequate model structure requires changes in distributional assumptions, model parameters, or covariates.

**Estimates of survival.**—Using MARK, we calculated maximum likelihood estimates of the multinomial parameters of survival and tag reporting based on an observed matrix of recaptures. Candidate models were fit to the tag recovery data and arranged in order of fit by the overdispersion-corrected second-order adjustment to QAIC\(_c\) (Akaike 1973; Burnham and Anderson 2002), namely,

\[
\text{QAIC}_c = -2 \cdot \log(\hat{\ell}(\hat{\beta})) + 2K + \frac{2K(K+1)}{(n-K-1)}.
\]

\(\log(\hat{\ell}(\hat{\beta}))\) is the log of the likelihood of the parameters (\( \hat{\beta} \)) given the data;

\[K = \text{the number of estimable parameters;}
\]

\[n = \text{the finite sample size;}
\]
Evidence ratios (i.e., the ratios of Akaike weights between two models) were calculated for comparing models (Burnham and Anderson 2002). Annual survival rates were estimated for fish 711 mm TL and larger. Annual survival was calculated as a weighted average across all models, where weight is a function of model fit (Buckland et al. 1997). Survival is inestimable for the terminal year in the time-saturated model, so the time-saturated model was excluded from the model-averaged survival estimate for the terminal year. A weighted average of unconditional variances (conditional on the set of models) was calculated for the model-averaged estimates of survival (Buckland et al. 1997).

Bias-adjusted estimates of survival.—Because we model dead recoveries, survival estimates were adjusted by annual estimates of live-release bias (Smith et al. 2000), namely,

\[
\text{bias} = - \left\{ \frac{\theta \cdot P_L f}{1 - (1 - \theta \cdot P_L f)} \right\};
\]

\[
\theta = 0.92 \quad \text{(based on an 8% hook-and-release mortality rate from Diodati and Richards 1996)};
\]

\[
P_L = \text{the annual proportion of tagged striped bass released alive};
\]

\[
f = \text{the annual recovery rate estimated with a Brownie recovery model (Brownie et al. 1985)};
\]

\[
\lambda = \text{the reporting rate}.
\]

Annual and geographic-based reporting rates are desirable, but unavailable; consequently, we used a constant reporting rate of 0.43 based on a Delaware Division of Fish and Wildlife Agency’s high-reward tag study (D. Kahn and C. Shirey, unpublished data).

Estimates of \( F \).—Instantaneous fishing mortality (\( F \)) was derived by converting the adjusted survival (\( S \)) to total mortality (\( Z \)) and then subtracting a constant value (\( M = 0.15 \)) for natural mortality, where \( F = -\log(S) - 0.15 \). When using a constant rate of natural mortality, a change in \( Z \) results in an equal change in \( F \). Uncertainty in each annual estimate of \( F \) (i.e., 95% confidence intervals) was derived from model-averaged unconditional variances of the adjusted survival estimates.

Results

Based on tag recovery locations (Figure 2), striped bass tagged off North Carolina and Virginia in winter migrated northward during summer and southward during fall (Table 2). During summer, tagged individuals were recovered in Maine, but the largest numbers were recovered from New York to Massachusetts, as well as waters of Maryland (Table 2). During spring months (April, May, and June), the largest numbers of tagged striped bass were caught within waters of Maryland (Chesapeake Bay) and New York (Hudson River). Chi-square tests did not reject (\( \alpha = 0.05, \text{df} = 4 \)) the null hypothesis of mixing of newly and previously tagged striped bass for the four periods of 1999–2001 (\( \chi^2 = 2.3, P = 0.68 \)), 2000–2002 (\( \chi^2 = 3.8, P = 0.44 \)), 2001–2003 (\( \chi^2 = 0.76, P = 0.94 \)), and 2002–2004 (\( \chi^2 = 5.7, P = 0.22 \)).

Regulatory period models highly influenced the model-averaged estimates of both survival and derived fishing mortality rates. Model-averaged estimates of annual survival and fishing mortality rates were derived primarily from four competing models. These top four competing models had QAIC\(_c\) values that changed by less than 2.2; the combined QAIC\(_c\) weight was 0.792, and all QAIC\(_c\) values were influenced by regulatory period parameterization (Table 3). Evidence ratios comparing the best approximating model to the next three competing models were 2.2, 2.7, and 2.9.

The analysis, however, yielded 10 of the 12 candidate models had QAIC\(_c\) values that changed by less than 7.0, indicating considerable model selection uncertainty (Table 3). Trend-based models, time-saturated models, and constant models for both survival and
Table 2.—Geographic and temporal distributions of the number of striped bass tag recoveries (1988–2004) by state and month; fish were tagged during 1988–2004.

<table>
<thead>
<tr>
<th>State</th>
<th>Jan</th>
<th>Feb</th>
<th>Mar</th>
<th>Apr</th>
<th>May</th>
<th>Jun</th>
<th>Jul</th>
<th>Aug</th>
<th>Sep</th>
<th>Oct</th>
<th>Nov</th>
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<td>Total</td>
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<td>53</td>
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<td>314</td>
<td>293</td>
<td>187</td>
<td>138</td>
<td>190</td>
<td>138</td>
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<td>1,737</td>
</tr>
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</table>

tag-reporting parameters were negligibly or not supported by the data; their evidence ratios were above 5 from comparisons with the best approximating model.

The model-averaged estimates of the bias-adjusted annual survival rates and proportion of fish released alive decreased during the 17-year time frame (Table 4). Estimates of bias-adjusted fishing mortality (range 0.00–0.26), which increased during the time series, were lowest during the restricted harvest period (1988–1989), intermediate during the interim fishery (1990–1994), and highest during the full fishery period (1995–2004; Table 4; Figure 3). Estimates of F from recent years were below target and threshold values of fishing mortality rate, but upper 95% confidence intervals (a measure of uncertainty) for recent F estimates exceeded these biological reference points (Table 4).

Goodness-of-fit statistics (GOF), $c$-estimated model fit and overdispersion, and annual recovery rates from a Brownie model (the proportion of live releases and live-release bias) were estimated for bias-adjustment of $F$. Low estimates of both GOF $P$-value ($P = 0.032$) and overdispersion ($c = 1.83$) were interpreted as an overdispersion contribution to lack of fit, which we addressed with $c$ adjustment (Burnham and Anderson 2002). Estimates of annual recovery rates ($f$) from a Brownie model were relatively constant across the 17-year time frame (range, 0.052–0.114; Table 4). The location and timing of tag recoveries reflect the spawning and foraging migration patterns of Atlantic coast striped bass as well as the distribution of fishing effort. The geographic and temporal distributions of tag recoveries were consistent with published patterns of striped bass migration, where northward movements in summer were followed by southward returns during fall (Boreman and Lewis 1987; Dorazio et al. 1994). The large numbers of springtime tag recoveries in the

### Discussion

After confirmation of typical north–south migration patterns and mixing of tagged cohorts, we estimated a 17-year time series of annual survival and fishing mortality rates for a southern-overwintering aggregate of migratory Atlantic striped bass. Our tag-based assessment of this mixed stock of striped bass yielded two important results. First, data supported regulatory-period models and indicated an association of temporal change between $F$ and harvest regulations. Second, estimates of $F$ increased across the 17-year time frame, but penultimate and terminal-year estimates of $F$ are below the management target and threshold values and do not indicate that Atlantic coast striped bass stocks are being overfished.

The location and timing of tag recoveries reflect the spawning and foraging migration patterns of Atlantic coast striped bass as well as the distribution of fishing effort. The geographic and temporal distributions of tag recoveries were consistent with published patterns of striped bass migration, where northward movements in summer were followed by southward returns during fall (Boreman and Lewis 1987; Dorazio et al. 1994). The large numbers of springtime tag recoveries in the

Table 3.—Model selection statistics used to derive model-averaged estimates of annual survival rates of striped bass ($\geq 711$ mm total length) tagged during the North Carolina winter trawl survey, 1988–2004. The term QAIC$_c$ is Akaike’s information criterion adjusted for small sample size and overdispersion. The third column shows the differences in the value of QAIC$_c$ between the best model (listed first) and each of the other models. See Table 1 for model descriptions; $k$ is the number of estimatable parameters in each model.

<table>
<thead>
<tr>
<th>Model</th>
<th>QAIC$_c$</th>
<th>AQAIC$_c$</th>
<th>QAIC$_c$ weights</th>
<th>$k$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$S(p) r(p)$</td>
<td>11,629.86</td>
<td>0.000</td>
<td>0.367</td>
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<td>$S(v) r(p)$</td>
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<td>1.989</td>
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<td>5.783</td>
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<td>$S(t) r(t)$</td>
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<td>0.001</td>
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<tr>
<td>$S(t) r(t)$</td>
<td>11,671.69</td>
<td>41.830</td>
<td>0.000</td>
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</table>
spawning areas of the Maryland portion of the Chesapeake Bay and the Hudson River probably represent spawners and further support the mixed-stock status of the overwintering aggregate off the coast of North Carolina. The geographic separation in peak harvest numbers between Massachusetts and Maryland during summer reflects fishing effort.

Incremental increases in the derived estimates of annual fishing mortality rates and decreases in the proportion of fish released alive coincided with periods of regulatory change during the 17-year time frame. The low estimates of \( F \) and high estimates of the proportion of fish released alive during 1988–1989 reflect the restricted harvest period (which included a moratorium in Chesapeake Bay, ASMFC 1990).

During 1990–1994, estimates of \( F \) and the proportion of fish released alive were intermediate relative to other values within the 17-year time series and reflect removal of the moratorium. Further, the highest values of \( F \) within the series occurred during 1995–2004, after managers declared a restored fishery and further liberalized harvest restrictions. Decreases in \( F \) consistent with regulatory periods were also found for striped bass tagged on the Maryland spawning grounds of the Chesapeake Bay and Hudson River, New York (Smith et al. 2000).

A large number of competing models and wide confidence intervals reflect uncertainty in estimates of \( F \). The data did not support a single best-approximating model, but rather supported four competing models all

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**Table 4.—Estimates of annual survival (\( S \)) and fishing mortality (\( F \)) rates of striped bass (≥711 mm total length), both unadjusted and adjusted for reporting rate (0.433), bias from live release, and hooking mortality rate (0.08); CI = confidence interval.**

<table>
<thead>
<tr>
<th>Year</th>
<th>Unadjusted</th>
<th>Live release</th>
<th>Adjusted</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>( S )</td>
<td>( F )</td>
<td>Recovery rate</td>
</tr>
<tr>
<td>1988</td>
<td>0.723</td>
<td>0.174</td>
<td>0.094</td>
</tr>
<tr>
<td>1989</td>
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**Figure 3.—Annual estimates of fishing mortality rates of striped bass (≥711 mm total length) as derived from tagging data gathered during the North Carolina winter trawl fishery. Error bars represent 95% confidence intervals.**
influenced by regulatory-period parameterization (with total QAICc weight of 0.792). Model averaging avoided selection of a single best model, and allowed for inference based on a weighted average of parameter estimates. Further, model-averaging incorporated model selection uncertainty into the $S$ and $F$ estimates and reduced the possibility of overestimating the precision of parameter estimates (Buckland et al. 1997; Burnham and Anderson 2002). Although regulatory-period models were weighted heavily, a model of constant survival (with reporting influenced by regulatory period) was also weighted by 0.125 and reflects additional uncertainty in our $S$ and $F$ estimates. Selection of a constant survival model does not mean no variation in survival across the time series but suggests little year-to-year variation in annual survival relative to information contained in the sample data (Burnham and Anderson 2002). In practice, selection of a constant survival model during analysis of a long-term time series likely reflects inadequate sample sizes, and our data of annual tag–release numbers were sparse in some years (see Table A.1). The 95% confidence intervals of 0.134–0.422 for penultimate-year estimates and 0.126–0.434 for terminal-year estimates ($F = 0.26$ for both) overlap and reflect uncertainty in relation to the current biological reference points of 0.41 (threshold $F$) and 0.30 (target $F$).

The proportion of live releases, tagging mortality, and tag loss can bias estimates of $F$. The downward trend in both the proportion of live releases and the associated live-release bias (i.e., catch-and-release bias) probably reflect regulatory changes in the fishery. Without adjustment for catch-and-release bias, we would have overestimated $F$ in recent years. The unadjusted $F$ estimates for penultimate ($F = 0.307$) and terminal ($F = 0.302$) years were at the target fishing mortality rate of 0.30 and would have elicited a different management response without correction for catch-and-release bias. Although unmeasured for this study, tagging mortality and tag loss will cause an underestimate of $F$. Estimates of tag-induced mortality, however, are low—0% (Goshorn et al. 1998) and 1.3% (Rugolo and Lange 1993)—and were excluded from bias adjustments. Additionally, we did not correct for tag loss given the low estimates of 0% (Goshorn et al. 1998), 2% (Dunning et al. 1987), and 2.6% (Sprankle et al. 1996).

The Cooperative Winter Tagging Cruise study and associated estimates of $F$ represent only one of eight tagging programs and sources of $F$ estimates within the ASMFC management process. Tagging data from this winter study support an increasing trend in $F$ but do not indicate that stocks are overfished. Our approach to tag-based estimates of $F$ represents current methods of the ASMFC Striped Bass Tagging Committee. Tag-based estimates of $F$ should be considered with other estimates of $F$ within the striped bass management process.

Acknowledgments

Members of the Atlantic States Marine Fisheries Commission’s striped bass committees have contributed to this work through discussions and advancement of tag-based assessments of migratory striped bass stocks along the Atlantic coast. We thank W. W. “Bill” Cole, Jr., for his many years of service as principal coordinator and chief scientist for the Cooperative Winter Tagging Cruises. Sara Winslow of the North Carolina Division of Marine Fisheries has served also as chief scientist and a Scientific Party member mainstay for almost every cruise. We thank all those staff of various agencies and universities along the east coast who have served as scientific party members for one or more cruises. The cruises would not be possible without the annual efforts of National Oceanic and Atmospheric Administration (NOAA) staff, commanding officers and crews of the NOAA research vessels Albatross IV, Chapman, and Oregon II. We thank also the commanding officer and crew of the National Science Foundation research vessel Cape Hatteras (which was employed in 2004), operated collaboratively by Duke University and the University of North Carolina-Chapel Hill. Funding for the Cooperative Winter Tagging Cruises was provided by the Atlantic States Marine Fisheries Commission, NOAA, and the U.S. Fish and Wildlife Service (Maryland Fisheries Resources Office, Annapolis, Maryland), and South Atlantic Fisheries Coordination Office (Morehead City and Raleigh, North Carolina). We especially thank Tina McCrobie and the Maryland Fisheries Resources Office of the U.S. Fish and Wildlife Service, and the Maryland Department of Natural Resources, Fisheries Division, for managing striped bass tagging data. The use of trade names does not imply government endorsement of commercial products.

References


ASMFC (Atlantic States Marine Fisheries Commission). 1990. Source document for the supplement to the striped


Appendix follows
### Appendix: Striped Bass Tag and Recovery Data

#### Table A.1.—Tag-recovery matrix of striped bass (≥711 mm total length) tagged off the coasts of North Carolina and Virginia from 1988 to 2004; \( N \) = number tagged each year.

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