Discerning patterns in patchy data: a categorical approach using gulf menhaden, *Brevoortia patronus*, bycatch

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Abstract.–We use two categorical techniques to explore patchy data on releasable bycatch in the U.S. Gulf of Mexico menhaden fishery. With no previous statistical studies on releasable bycatch, this paper serves both as an analysis of patchy data with categorical techniques and as a quantitative description of a biologically important portion of the menhaden bycatch.

By means of exploratory analyses with loglinear and logit models, we determined spatial and temporal patterns in bycatch of the menhaden fishery. Contrasts revealed that at fishing grounds east of the Mississippi River, the probability of observing fishing sets with high bycatch rates in spring and summer were greater than in the fall. Furthermore, spring bycatch rates were higher in fishing areas east of 89°W than in fishing areas west of 93°W.

Correspondence analysis indicated that the fate of the releasable bycatch could be classified into three major groupings. The first group, species associated with being gilled, was composed primarily of *Micropogonias undulatus, Trichiurus lepturus, Chloroscombrus chrysurus,* and *Cynoscion arenarius.* The second group consisted of species associated with being dead or disoriented; it included the requiem sharks *Caranx hippos* and *Sciaenops ocellatus.* The third group included those fish that were associated with being put into the hold, those kept by the crew, or those whose fate was unknown. These included *Arius felis, Bagre marinus,* and *Scomberomorus maculatus.*

Seasonal and spatial associations of bycatch species were also examined with correspondence analysis. From April through August, two distinct bycatch species assemblages were observed that separated the fishery at a longitude of 91°W. From September through October, a shift in the species assemblage indicated that the western region of the fishery (west of 93°W) appeared to have a assemblage distinct from the rest of the fishery.

Manuscript accepted 18 August 1997.
Fishery Bulletin 96: 193–209 (1998).

Statistical analyses of fisheries data often require consideration of the patchy distribution of aquatic organisms (e.g. Andrew and Pepperell, 1992). This is especially true with standard regression and ANOVA techniques because they rely on the assumptions of normality of residuals and homogeneous variances. Often, transformations of the response variable can be used to satisfy these assumptions (Underwood, 1981). However, in many instances suitable transformations cannot be found. In such cases, the use of standard regression and ANOVA techniques will result in gross violations of the model assumptions and affect the results of the analysis (Underwood, 1981). Where suitable transformations cannot be found, nonparametric techniques such as the Kruskal-Wallis one-way ANOVA or ANOVA on the ranked data have been used. As Underwood (1981) has noted, one of the implied assumptions of these methods, that of equal variances among treatments, is usually not satisfied when dealing with patchy data. A further problem with such techniques is that one cannot explore significant differences with the use of contrasts or multiple comparison tests, except with the Kruskal-Wallis one-way ANOVA. Therefore, when using these techniques, one is limited to the use of empirical methods to examine significant differences.

Studies on fisheries bycatch have used a variety of statistical methods. Andrew et al. (1995) used ANOVA with suitable transformations to evaluate the bycatch in an Australian stow-net fishery for school prawns, and Austin et al. (1994) used the Kruskal-Wallis one-way ANOVA to examine geographical differences in the bycatch of the Atlantic menhaden fishery. Hudson (1990), examining the shrimp and fish bycatch assemblages of the Canadian Eastern Arctic, used an intermediate linkage clustering algorithm to examine abundance patterns and species associations among different regions. More unique solutions have been presented by Richards et al. (1994), who proposed the use of a modified generalized logit model to carry out a categorical form of response surface analysis, allowing the estimation of transformation parameters on the explanatory variables along with other parameters. Perkins and Edwards (1996) have used mixture models, consisting of the negative binomial distribution with added zeros, as a solution to analyzing bycatch with many zero observations.

In this paper we use loglinear and logit models to examine patchy data with bycatch from the gulf menhaden, *Brevoortia patronus,* fishery. Menhaden bycatch can be classified into two groups: 1) bycatch that is pumped directly into the hold with the menhaden and that is termed "automatically retained bycatch"; and 2) all other bycatch that is termed "releasable bycatch." In this paper we analyze only releasable bycatch. With the exception of Christmas et al. (1960) and Condrey,[1] previous studies on menhaden bycatch have not taken releasable bycatch into account. Furthermore, all previous work has been qualitative in nature. Our examination of releasable bycatch, however, serves both as an analysis of patchy data with categorical techniques and as a quantitative description of a biologically important portion of the menhaden bycatch.

Species taken as bycatch may be caught because they are associated with the target species or simply because they were encountered on a random basis (Hall, 1996). An analysis of the structure of bycatch species assemblages associated with a fishery can provide valuable information for both management and ecological purposes. Hudson (1990), examining shrimp and fish bycatch assemblages in the Canadian Eastern Arctic, observed three associations that she proposed were related to the origin of the predominant water masses in the region. Harris and Poiner (1991) documented changes in the species composition of demersal fish fauna over a 20-year period in the Gulf of Carpentaria, Australia, and suggested that increases in the benthopelagic taxa over this period could be partially explained by discard of bycatch in the banana prawn fishery. Because little information on spatial and temporal associations for species assemblages of the menhaden fishery exists, we describe areal and temporal associations for these species, together with the fate of these species, as a first step in accruing such information. We have used correspondence analysis, a categorical form of ordination, to describe these association patterns.

The fishery

The U.S. Gulf of Mexico menhaden fishery has existed since the late 1800's (Nicholson, 1978) and is the largest fishery in tonnage in the northwestern Gulf of Mexico. Estimates of menhaden landings for the 1994 fishing season were approximately 0.7 million metric tons (t) (Leard et al., 1995). Although gulf menhaden is the primary clupeid sought, finescale menhaden, *B. gunteri*, yellowfin menhaden, *B. smithi*, and Atlantic thread herring, *Opisthonema oglinum,* are occasionally taken opportunistically (Leard et al., 1995).

Schools of menhaden are located visually by the senior crew aboard the fishing vessel or with the help of spotter planes. A purse seine, deployed from a pair of small (12.2-m) purse-seine boats, is used to encircle the school. Once the school is encircled, the bottom of the net is drawn up to hold the catch. The seine is then retrieved mechanically by each purse-seine boat until the fish are confined to a small section of the net. The catch is then pumped into the refrigerated hold of a larger (43 to 61 m) carrier vessel. The number of times the purse net is set each day depends on the availability and size of schools. Schools contain from 3 to 100 t of menhaden (Leard et al., 1995). Once the hold of a vessel is full, or a trip is otherwise complete, the menhaden are transported to one of the processing plants located from Moss Point, MS, to Cameron, LA. Although the fishing area extends from Apalachicola, FL, to Freeport, TX, more than 86% of the menhaden caught from 1990 to 1994 were taken off Louisiana (Leard et al., 1995).

Materials and methods

Loglinear and logit models

As reviewed in Agresti (1990) and Freeman (1987), loglinear and logit models are special cases of the generalized linear models introduced by Nelder and Wedderburn (1972). Agresti (1990) summarizes a generalized linear model as "a linear model for a transformed mean of a variable having a distribution in the natural exponential family."

Loglinear models describe association patterns among categorical variables. With this approach, cell counts in a contingency table are modeled in terms of association among the variables. Loglinear models may be viewed as analogous to correlation analysis where cell counts in a loglinear contingency table are treated as independent Poisson variables.

In a $I \times J$ table, where $N = IJ$ cells consisting of n multinomial samples, let n_k denote the count of the

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[1] Condrey, R. E. 1994. Bycatch in the U.S. Gulf of Mexico menhaden fishery. Results of onboard sampling conducted in the 1992 fishing season. Coastal Fisheries Institute, Louisiana State Univ., Baton Rouge, LA. Final report for NA27F0007-01, 42 p. [Available through the NOAA library, Silver Springs, MD.]

k^{th} cell, and let $m_k = E(n_k)$ represent the expected value where $k = 1.....N$. The probabilities $\{\pi_{ij}\}$ for that multinomial distribution form the joint distribution of two categorical responses. These two responses are statistically independent when $\pi_{ij} = \pi_{i+}\,\pi_{+j}$ for all i and j.

If there is a dependence between the two variables, then all expected values of each cell (m_{ij}) are > 0. The loglinear model for this two-way table can be written as

Table 1

Hypothetical example used to explain odds ratios and conditional probabilities. $\pi_{j1|i}$ is the conditional probability of observing bycatch given a particular area. $\pi_{j2|i}$ is the conditional probability of observing no bycatch given a particular area.

	Bycatch			Conditional probabilities	
	Presence (j_1)	Absence (j_2)	Total sets	$\pi_{j1\mid i}$	$\pi_{j2\mid i}$
Area A (i_1)	4	6	10	0.40	0.60
Area B (i_2)	2	4	6	0.33	0.67
Total	6	10	16		

$$\log m_{ij} = \mu + \lambda_i^x + \lambda_j^y + \lambda_{ij}^{xy},$$

$$\begin{aligned} \text{where } \mu &= \Sigma_i \Sigma_j \log m_{ij} / IJ; \\ \lambda_i^x &= \Sigma_j \log m_{ij} / J - \mu; \\ \lambda_j^y &= \Sigma_i \log m_{ij} / I - \mu; \text{ and} \\ \lambda_{ij}^{xy} &= \log m_{ij} - \lambda_i^x - \lambda_j^y + \mu. \end{aligned}$$

This model perfectly describes any set of positive expected frequencies and is referred to as the saturated model. The right-hand side of this equation resembles the formula for the cell-means ANOVA. The parameters $\{\lambda_i^x\}$ and $\{\lambda_j^y\}$ are deviations about a mean and $\Sigma_i \lambda_{ij}^{xy} = \Sigma_j \lambda_{ij}^{xy} = \Sigma_i \lambda_i^x = \Sigma_j \lambda_j^y = 0$. This model can also be described in the notation form as $[XY]$.

A saturated loglinear model always expresses a given table of categorical data perfectly. This model has the maximum achievable log likelihood because it is the most general model, with as many parameters as observations. However, it is possible that a simpler model may provide a fit as statistically good as that of the saturated model. How well this model fits is represented by the scaled deviance, a function of twice the difference in the log likelihoods of the saturated model and the simpler model. In addition to testing the fit of a model, one can use the deviance to diagnose lack of fit through residual analysis.

For example, consider a three-dimensional saturated model with variables X, Y, and Z. For this model $[XYZ]$, $\log m_{ijk} = \mu + \lambda_i^x + \lambda_j^y + \lambda_k^z + \lambda_{ij}^{xy} + \lambda_{ik}^{xz} + \lambda_{jk}^{yz} + \lambda_{ijk}^{xyz}$. When $\lambda_{ijk}^{xyz} = 0$, there is no three-factor interaction, and the association between two variables is identical at each level of the third variable and reduces to the loglinear model $[XY\,XZ\,YZ]$. Further, if $\lambda_{ijk}^{xyz} = 0$ and $\lambda_{jk}^{yz} = 0$, then for any given level of X, Y and Z are conditionally independent $[XY\,XZ]$. Similarly if $\lambda_{ijk}^{xyz} = 0$, $\lambda_{jk}^{yz} = 0$ and $\lambda_{ik}^{xz} = 0$, then Z is jointly independent of X and Y $[XY\,Z]$. Finally if $\lambda_{ijk}^{xyz} = 0$, $\lambda_{jk}^{yz} = 0$, $\lambda_{ik}^{xz} = 0$, and $\lambda_{ij}^{xy} = 0$, then X, Y, and Z are mutually independent $[X\,Y\,Z]$. With these criteria and beginning with a saturated model, we used a stepwise model selection procedure with deviance in the form of the G^2 test statistic to find a simpler model that fits as well as the saturated model. This simpler model would enable one to explore multidimensional tables to find simpler representations of the information contained therein.

Another advantage of loglinear models is that when one of the variables can be modeled as a response, and the others as explanatory variables, certain loglinear models are equivalent to logit models with categorical explanatory variables. Such logit models enable us to study the problem of interest in a manner analogous to ANOVA.

Many categorical response variables have only two categories. The response can be classified either as a success or a failure. The Bernoulli distribution, which belongs to the natural exponential family, forms the basis of modeling the logit model. For such a dichotomous variable, the probability of observing response 0 can be defined as $P(Y{=}0) = \pi$, and the probability of observing response 1, as $P(Y{=}1) = 1 - \pi$. The link function for this model, $\log \pi_i \,/\, (1-\pi_i)$, known as the logit, is equivalent to the log odds.

Consider the following example, where we examine the presence or absence of bycatch in two areas. In this example (Table 1), the 2 × 2 table has rows i_1*(area* 1), and i_2 (*area* 2) and columns j_1 (*presence of bycatch*) and j_2 (*absence of bycatch*). The counts in the cells of the table are the number of units of effort (individual sets) observed in each category.

In this case, the odds, Ω_{i1}, of observing j_1 (*presence of bycatch*) given you are in category i_1 (*area* 1) is computed as the ratio of the conditional probabilities of observing a set with bycatch to that of observing a set with no bycatch in area 1:

$$\left\{\pi_{j1|i1}/\pi_{j2|i1}\right\} \text{ is } 0.4/0.6 = 0.67.$$

Similarly, the odds, Ω_{i2}, of observing j_1*(presence of bycatch)* given you are in category i_2 (*area* 2) is computed as the ratio of the conditional probabilities of observing a set with bycatch to that of observing a set with no bycatch in area 2:

$$\{\pi_{j1|i2}/\pi_{j2|i2}\} \text{ is } 0.33/0.67 = 0.5.$$

The odds ratio, θ, is computed as

$$\Omega_{i1}/\Omega_{i2} \text{ is } 0.67/0.5 = 1.34.$$

Thus, the odds of observing response j_1 *(presence of bycatch)* is 1.34 times more likely for row i_1 *(area 1)* than for row i_2 *(area 2)*. An odds ratio of 1 indicates that you are equally likely to observe response j_1*(presence of bycatch)* for row i_1 *(area 1)* and row i_2 *(area 2)* and thus indicates independence between the rows and columns of the table.

The logit model has two forms. One form occurs where the explanatory variables are continuous and is the logistic regression model. The second occurs where the explanatory variables are categorical. The logistic regression model is analogous to a regression model, whereas the second type is analogous to an ANOVA model.

For the previous example, a model with a single categorical explanatory factor (area), the logit form of the model is

$$\log(\pi_{j1|i}/\pi_{j2|i}) = \alpha + \beta_i^{Area},$$

where α = the mean of the logits; and
β_i^{Area} = the deviation from the mean for row i.

β_i describes the effects of the factor on the response. For this model the higher β_i becomes, the higher the logit in row i, and the higher the value of $\pi_{j1|i}$. The constraints on this model are $\Sigma\ \beta_i = 0$. In this case the right-hand side of the equation resembles the cell-means model of a one-way ANOVA. This logit model would be equivalent to $\log(m_{ij1}) - \log(m_{ij2}) = 2\lambda_i^{Area} + 2\lambda_{j1}^{Bycatch} + 2\lambda_{ij1}^{AreaBycatch}$ in loglinear form.

Bycatch sampling and data set description

Bycatch from the gulf menhaden fishery was sampled April through October 1995 by two to three onboard samplers on a total of twenty-seven week-long trips aboard vessels operating from menhaden processing plants in the U.S. Gulf of Mexico. To maximize coverage of the Gulf, samplers boarded vessels from ports in the western, central, and eastern regions in a given week as often as possible. During each sampling trip, all sets made by the vessel were alternatively sampled, either for releasable bycatch or automatically retained bycatch. For all sets sampled, the presence of dolphins in the vicinity was also noted by the observers. In addition, the boat captains visually estimated catch in standard menhaden (1,000 standard menhaden [~305 kg]) and recorded the latitude and longitude of a set location. The location was used to identify in which National Marine Fisheries Service (NMFS) statistical zone (Fig. 1) the set was made (after Kutkuhn, 1962).

To collect releasable bycatch data, samplers observed the purse seine from the time it was brought alongside the carrier ship and throughout the pumping procedure, until the net was emptied and cleaned. During this time, the species, number, and fate of the releasable bycatch were recorded. The seven categories of bycatch fate were as follows: gilled in the net (gilled); kept by the crew for consumption (kept); released with no apparent harm (released healthy); released seriously injured or dead (released dead); released after being bruised or after being kept in the set for a long time (released disoriented); collected by the crew from the net or deck and put into the hold (caught and put in hold); and observed in the net but fate unknown (unknown).

Statistical analysis

Preliminary analysis For the variables bycatch number, bycatch percentage, and estimated catch, we calculated a series of commonly used statistical descriptors, namely the mean, standard deviation, 95% confidence intervals, median, skewness, and kurtosis. In addition, we also calculated the winsorized mean and its standard deviation.

We initially attempted to examine spatial and temporal patterns in the bycatch with a two-way ANOVA model. For the analysis, data were classified into season (S) consisting of three groups: 1) spring (April through June), 2) summer (July through August), and 3) fall (September through October). Adjacent NMFS zones (Fig. 1) were combined to form four area (A) groups: 11–12, 13–14, 15–16, and 17–18.

Bycatch patterns were examined with two response variables: 1) bycatch numbers; and 2) bycatch percentage ([bycatch number/total catch] × 100). For each of these two response variables, spatial and temporal patterns were examined by using the ANOVA model with season, area, and their interaction term as independent variables. Because we anticipated that neither model would satisfy the model assumptions of normality of residuals and homogeneous variances, we also examined the models by using the log and square-root transformations for both response variables. In addition we also used 2 arcsin $\sqrt{}$ (bycatch/menhaden catch) suggested by Neter et al. (1990) for transformation of proportions. All seven models were examined to determine if model assumptions were met.

Spatial and temporal patterns in bycatch For our analysis using loglinear and logit models with categorical explanatory variables, we used a four-way contingency table with a unit of effort (the set) as the count. Our main interest was 1) to examine the spatial and temporal patterns in bycatch and 2) to determine if the presence of dolphins in the vicinity when the set was made might be an indicator of bycatch patterns.

Exploratory analysis with loglinear and logit models To examine bycatch as a response of interest with categorical models, a new dichotomous categorical variable, bycatch, based on the median bycatch percentage, was created. Each set was classified either as high bycatch if the bycatch rate of the set was greater than the median value of all sets or as low bycatch if the bycatch rate of the set was less than or equal to the median bycatch of all the sets. We used the median rate because it is a robust measure of central tendency. In deciding on possible criteria for defining this variable, more extreme conditions, such as bycatch rates greater than the 75th percentile, were considered. However, by choosing more extreme values, we increased the number of sparse cells and thus affected the validity of the G^2 test statistic.

In analyzing contingency tables, it is necessary that the number of cell counts with zero frequencies be low (a minimum expected value of 1 is satisfactory as long as <20% of cells have counts of 5 or less) for the test statistic to be valid (Agresti, 1990). To reduce the number of cells with zero frequencies, months and zones were combined, generating two new variables, season and area, corresponding to those used in the ANOVA. The presence of dolphins was used as a dichotomous variable, dolphins (*D*).

To identify the most appropriate and simplest loglinear model for the data using the variables season, area, bycatch, and dolphins, we employed a stepwise backward solution procedure commencing with the saturated loglinear model (Agresti, 1990). Here the saturated model

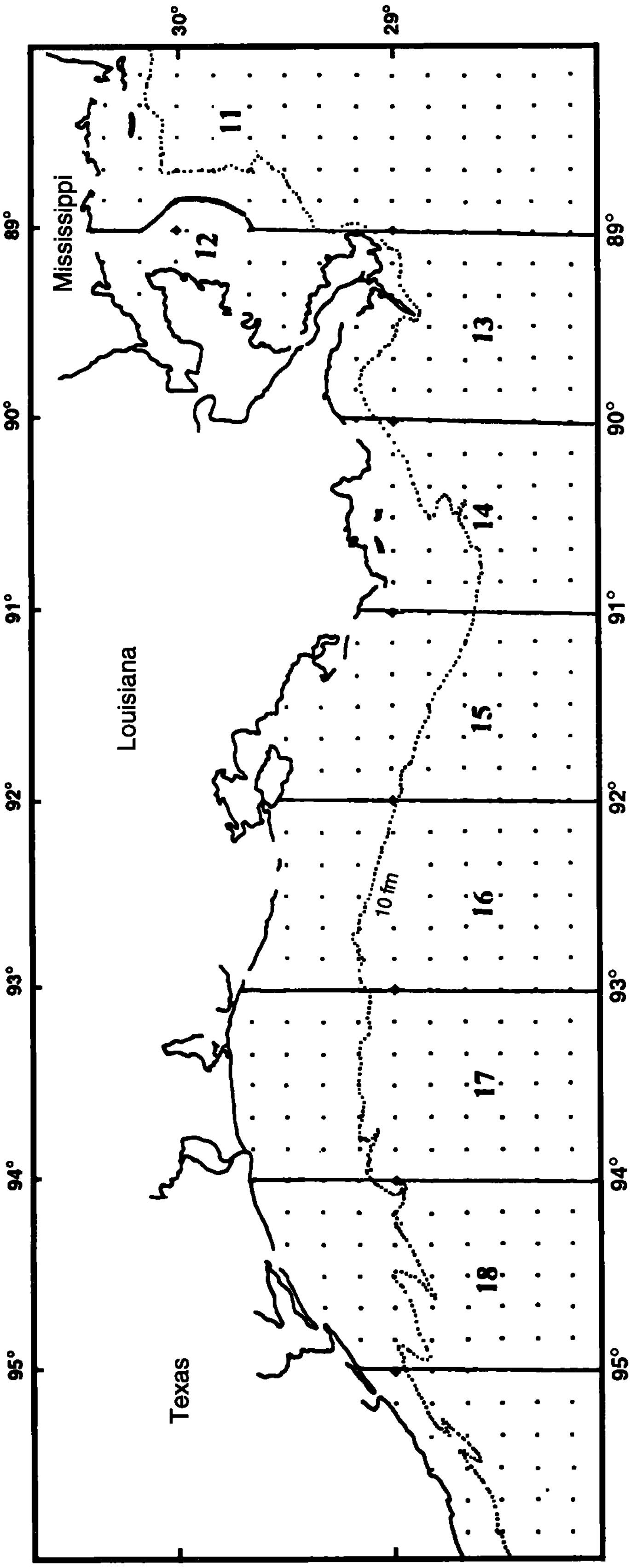

Figure 1

Map encompassing the extent of the U.S. gulf menhaden fishery from the Texas to Alabama coasts. The eight fishing zones numbered 11–18 are from NMFS (after Kutkuhn, 1962).

is denoted as [*SABD*] were *S* stands for season, *A* for area, *B* for bycatch, and *D* for dolphins, and the model includes all possible interactions up to and including the four-way interaction. Because the saturated model would naturally provide the best fit, we were interested in determining if a simpler model could be found that would also satisfy the criteria of a logit model with bycatch as the response variable. The standardized residuals of the resulting model were then examined to ensure that lack of fit was not a problem.

Contrasts We anticipated that we would have significant interaction terms in our analyses that would require detailed examination of interactions. For the logit form of the selected model, we constructed a series of contrasts that might help to explain the nature of these potential interactions. The contrasts of interest had two general forms:

1 Given a specific area, are the odds of observing a set with high bycatch the same between any two different seasons. This results in three unique contrasts for each area (spring vs. summer, spring vs. fall, summer vs. fall) and a total of 12 contrasts.

Let F_{ij} be the logit of high bycatch for season i and area j and let F_{hj} be the logit of high bycatch for season h and area j.

The hypotheses being tested were

$$H_0: F_{ij} - F_{hj} = 0,$$

where $F_{ij} = \alpha + \beta_i^S + \beta_j^A + \beta_{ij}^{SA}$;
$F_{hj} = \alpha + \beta_h^S + \beta_j^A + \beta_{hj}^{SA}$; and
h and i = spring, summer, and fall, such that $h \neq i$ for each j, and j = area 11–12, area 13–14, area 15–16, and area 17–18.

2 Given a specific season, are the odds of observing a set with high bycatch the same between any two different areas. This results in six unique contrasts for a given season and a total of 18 contrasts (11–12 vs. 13–14, 11–12 vs. 15–16, 11–12 vs. 17–18, ... ,15–16 vs. 17–18).

Let F_{ij} be the logit of high bycatch for season i and area j and let F_{ik} be the logit of high bycatch for season i and area k.

The hypotheses being tested were

$$H_0: F_{ij} - F_{ik} = 0,$$

where $F_{ij} = \alpha + \beta_i^S + \beta_j^A + \beta_{ij}^{SA}$;
$F_{ik} = \alpha + \beta_i^S + \beta_k^A + \beta_{ik}^{SA}$; and
j and k = area 11–12, area 13–14, area 15–16, and area 17–18, such that $j \neq k$ for each i, where i = spring, summer, and fall.

Because 30 contrasts were performed, the type-I error level of 0.10 was adjusted by using the Bonferroni technique, and only *P*-values less than 0.0033 were considered significant. The estimated odds ratios for the conditions associated with the hypotheses were calculated from the parameter estimates given by the analysis.

Bycatch species associations To examine the association between species and fates of the releasable bycatch, we used correspondence analysis on a species-by-fate table for all seasons and areas combined. Area and species associations of the releasable bycatch were also examined for each of the three seasons with correspondence analysis on species-by-area tables.

For all correspondence analyses we defined two groups of species. The first group, consisting of those species that were common in terms of number and occurrence, was used in the main table. Releasable bycatch species falling into this group had a minimum of 230 individuals and were found in at least 30% of the sets. The second group of species consisted of releasable bycatch that were less common; these were species for which a minimum of 30 individuals were observed, which occurred in at least 4% of the sets, and which did not meet our criteria for well represented species. These species were included as supplementary variables in our analysis (Greenacre, 1984). Supplementary variables are represented as points in the joint row and column space but are not used in determining the locations of the active rows and columns of the table. Species included in the main and supplementary table accounted for 97% of the total number of organisms observed during the study period. Species that did not meet these criteria were not used in the analyses.

Results

Preliminary analysis

A total of 15,579 bycatch organisms representing 62 species or species groups were observed as releasable bycatch in 257 sets. The estimated catch of standard menhaden per set ranged from 5,000 to 500,000 with a median, mean, and standard deviation of 50,000, 67,000, and 61,000 respectively. Skewness and kurtosis values of 2.5 and 10.7 indicated that the distribution of the estimated menhaden catch was positively skewed (Fig. 2).

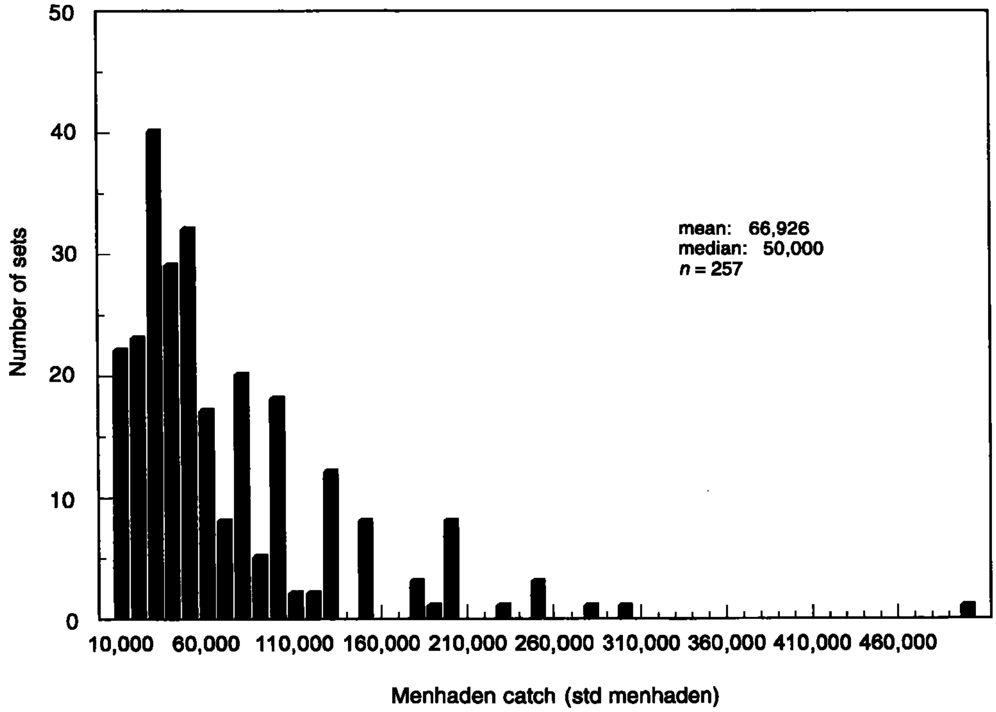

Figure 2

Distribution of menhaden catch sampled during the 1995 gulf menhaden fishing season. std = standard (1,000 standard menhaden ~ 305 kg).

The number of bycatch observed in each set ranged from 0 to 1,600 organisms, with a median, mean, and standard deviation of 15, 61, and 153 respectively. The winsorized mean and its standard deviation values were 53 and 6.9 respectively. The 95% confidence interval of the mean was between 41.8 and 79.3. The distribution of bycatch organisms was strongly positively skewed (5.9) and peaked sharply with a Kurtosis value of 47.2 (Fig. 3).

The bycatch percentage ranged from 0% to 4% with a median, mean, and standard deviation of 0.033%, 0.168%, and 0.48 respectively. The winsorized mean bycatch percentage and its standard deviation were 0.14% and 0.02, respectively. The 95% confidence interval of the mean was between 0.11% and 0.22%. The distribution of the bycatch percentage was also found to be positively skewed and strongly peaked, with skewness and kurtosis values of 5.4 and 32.7 respectively (Fig. 4).

Analysis of variance with bycatch, bycatch percentage, and their respective transformations, together with the arcsine transformation, did not meet model assumptions. In all cases the modified Levene's test indicated that the variances were nonhomogeneous, and both the residual plots and Shapiro-Wilk test indicated that the assumption of normality of residuals were not met. For example, for the response log (bycatch percentage + 1), the residuals of the model were not normally distributed (Shapiro-Wilk W=0.678, $P<W$=0.0001), and the residual plot indicated nonhomogenous variances. Furthermore, a modified Levine's test indicated that the variances were nonhomogeneous (F=6.21, df=11, df-error=245, $P>F$=0.0001). These characteristics suggest that the model assumptions were grossly violated and that ANOVA may not be an appropriate form of analysis in this case.

Spatial and temporal patterns in bycatch

Exploratory analysis using loglinear models With the backward selection procedure, loglinear models [SAB SAD DB] and [SAB SAD] (as defined in Table 2) satisfied the criteria for a logit model and had good fit (Table 2). The simplest of these models, [SAB SAD], was selected; this loglinear model corresponds to the logit model with categorical explanatory variables of the form

$$\log\frac{\pi_{highlik}}{\pi_{lowlik}} = \alpha + \beta_i^A + \beta_k^S + \beta_{ik}^{AS}, \qquad (1)$$

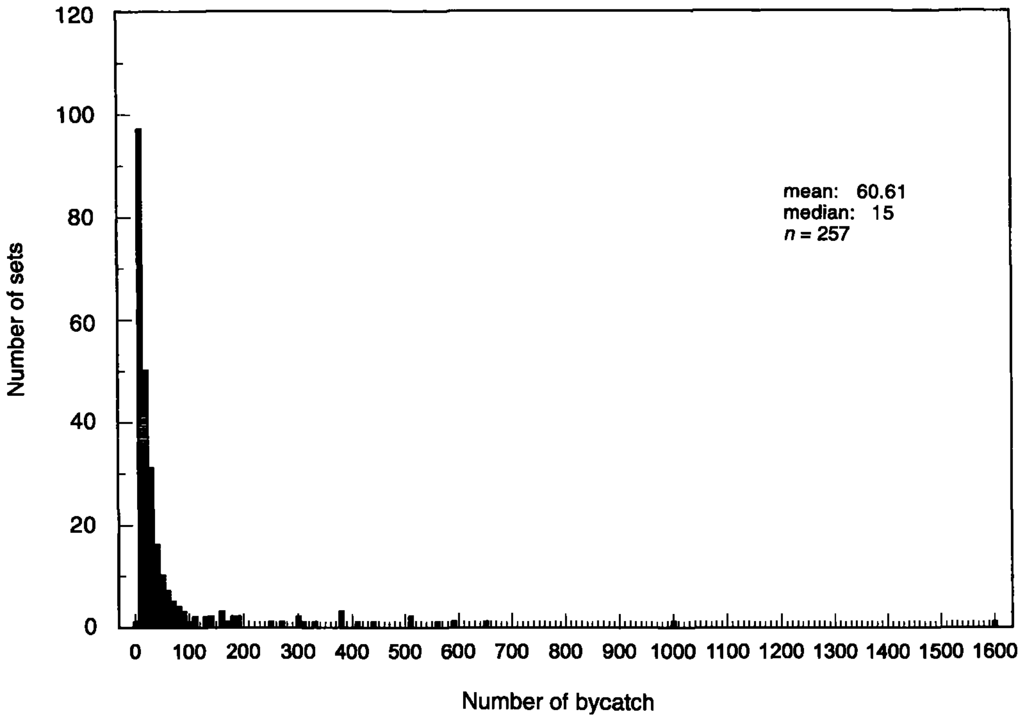

Figure 3

Distribution of bycatch sampled during the 1995 gulf menhaden fishing season.

Table 2

Summary statistics of loglinear models examined through the stepwise selection procedure. Models with fits as good as the saturated model are marked ns. (A=area, B=bycatch, D=dolphins, S=season). Models are presented hierarchically from most complex to simplest.

Loglinear model	df	G^2	*P*-value
Model without the four-way interaction term			
ADB SDB SAB SAD	6	11.64	0.0706
Selected models with 3 three-way interaction terms			
ADB SDB SAB	12	26.58	0.0089
ADB SDB SAD	12	27.62	0.0063
SDB SAB SAD^{ns}	9	13.41	0.1449
SAB SAD ADB	14	27.90	0.0147
Selected models with 2 three-way and 1 two-way interaction terms			
SDB SAB DA	15	29.38	0.0144
SAB SAD DB^{ns}	11	15.65	0.1547
Selected models with 2 three-way interaction terms			
SAB SAD^{ns}	12	17.82	0.1213
SDB SAB	18	45.83	0.0003
Selected models with 1 three-way and 2 two-way interaction terms			
SAB SD AD	18	34.89	0.0097
SAD SB AB	18	34.30	0.0116

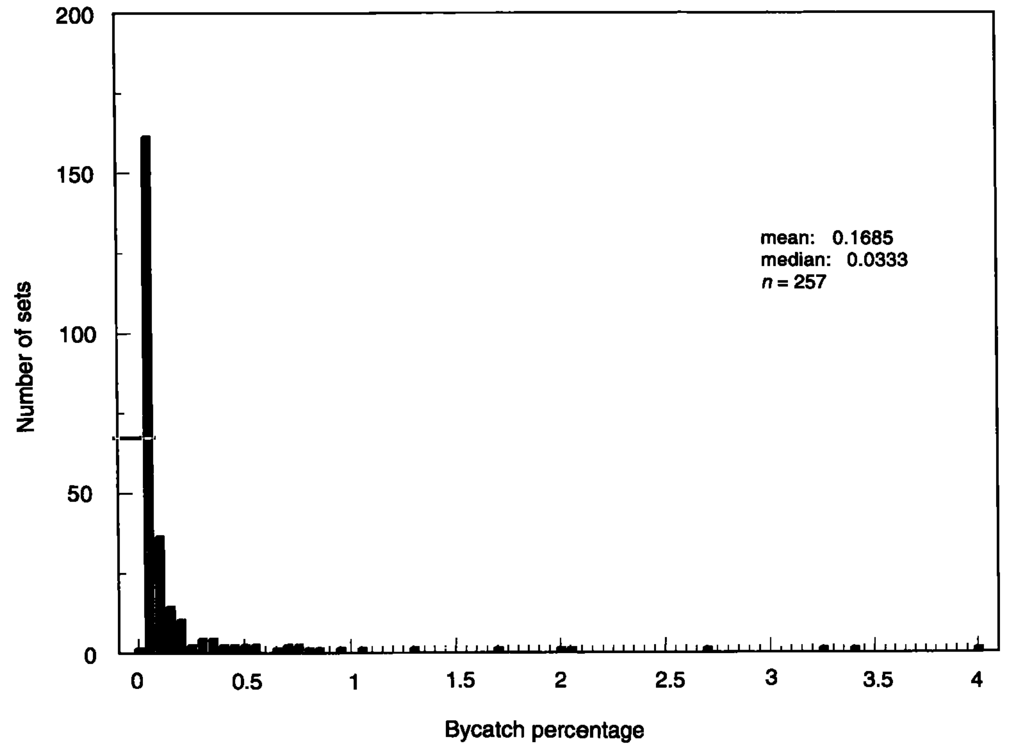

Figure 4

Distribution of the bycatch percentage in fishing sets sampled during the 1995 gulf menhaden fishing season.

where $\pi_{high\,|\,ik}$ = the probability of observing a set with high bycatch given area i and season k; and

$\pi_{low\,|\,ik}$ = the probability of observing a set with low bycatch given area i and season k.

This model, in loglinear and logit forms, had a G^2=17.82, with 12 df and a P-value=0.1213. Examination of the standardized residuals of the logit form revealed that none of the residuals had an absolute value greater than 1.45 and thus showed no evidence of lack of fit.

For the logit model, there was a significant interaction between season and area (Wald chi-square (χ^2)=14.65, df=6, P=0.0232). This interaction is reflected in the plot of probabilities of high (greater than 0.033%) bycatch for the area-season combinations (Fig. 5).

Table 3

Estimated odds ratios and Wald chi-square (χ^2) values of contrasts of observing high bycatch between seasons, given a set was made in a certain area. For example, the odds of observing high bycatch is 11.761 times greater during the spring season than during the fall season for area 11–12. * indicates a ratio significantly different from 1. Wald χ^2 values are in parenthesis. For all contrasts, Wald χ^2 df = 1.

Season	Area	Spring	Fall
Fall	11–12	11.761* (10.03)	
Summer	11–12	1.142 (0.05)	0.097* (9.25)
Fall	13–14	4.444 (2.24)	
Summer	13–14	0.555 (0.19)	0.125 (2.41)
Fall	15–16	0.699 (0.48)	
Summer	15–16	0.414 (4.29)	0.592 (0.97)
Fall	17–18	0.312 (2.11)	
Summer	17–18	0.218 (2.47)	0.899 (0.02)

Contrasts The contrasts of the odds of high bycatch between seasons for a given area indicate that the hypothesis of spring and summer seasons being the same for all four areas could not be rejected. Wald χ^2 values are presented in Table 3. The contrasts of the spring and fall seasons suggest that the only significant differences between these two seasons existed for area 11–12 (Wald χ^2=10.03, df=1, P>χ^2=0.0015). When the summer and fall seasons were contrasted, a significant difference between seasons was observed only for area 11–12 (Wald χ^2=9.25, df=1, P>χ^2=0.0024).

For contrasts of the odds of high bycatch between areas for a given season, significant differences between areas 11–12 and 17–18 in the spring season were observed (Wald χ^2=9.29, df=1, P>χ^2=0.0023). All other contrasts between areas for the spring, summer, and fall seasons were not significant. Wald χ^2 values for these contrasts are presented in Table 4.

The estimated odds ratios for the conditions associated with the hypotheses are also given in Tables 3 and 4. Only estimated odds ratios of contrasts for the rejected hypotheses of no difference are examined in detail.

If a fishing boat was in area 11–12, the odds of observing a set with high bycatch are about 11.8 times higher if a boat was fishing in the spring rather than the fall. Also, if a fishing boat was in area 11–12, the odds of observing a set with high bycatch are about 0.1 times (or approximately 10.3 times lower) in the fall than in the summer. It appears that for area 11–12, the odds of observing a set with high bycatch in the fall are significantly lower than in the spring or summer.

The third significant contrast indicated that for a vessel fishing in the spring, the odds of observing a set with high bycatch are about 10.8 times higher in area 11–12 than in area 17–18.

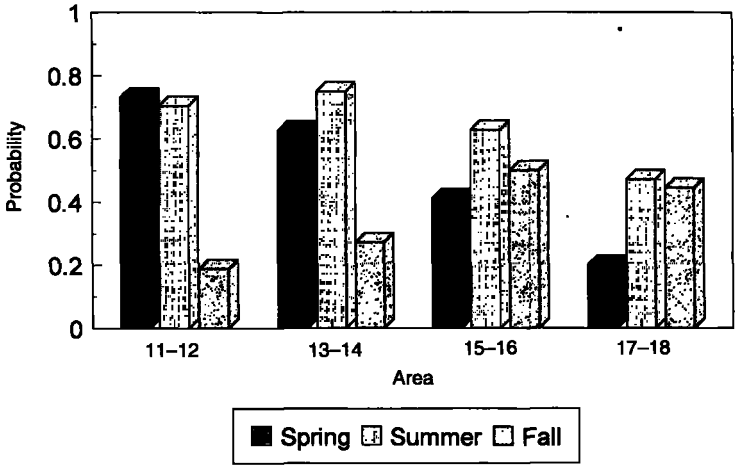

Figure 5

Probabilities of observing a set with high bycatch (> 0.033%) for the season-area combinations.

Table 4

Estimated odds ratios and Wald chi-square (χ^2) values for contrasts (in parentheses) for observing high bycatch between areas, given a set was made in a certain season. For example, the odds of observing high bycatch are 1.628 times greater in area 11–12 than in area 13–14. * indicates a ratio significantly different from 1. Wald χ^2 values are in parenthesis. For all contrasts, χ^2 df=1.

		Area		
Season	Area	11–12	13–14	15–16
Spring				
	13–14	1.628 (0.33)		
	15–16	3.877 (6.64)	2.380 (1.23)	
	17–18	10.857* (9.29)	6.666 (3.79)	2.799 (2.13)
Summer				
	13–14	0.791 (0.04)		
	15–16	1.407 (0.42)	1.777 (0.23)	
	17–18	2.671 (2.33)	3.374 (0.94)	1.898 (1.22)
Fall				
	13–14	0.615 (0.27)		
	15–16	0.230 (3.63)	0.375 (1.50)	
	17–18	0.288 (2.43)	0.468 (0.84)	1.250 (0.12)

Refining the final model We were interested, on the basis of our contrasts, in determining if a model with simpler dichotomous classes for areas and seasons would provide as good a statistical fit as this "full model" (Eq. 1). We compared three potential models that had one or both of these variables with reduced classes against the full model (Table 5). For these models we classified area into two groups: 1) east of the Mississippi River; and 2) west of the Mississippi River. Season was also classified into two groups; 1) early—sets sampled April through August; and 2) late—sets sampled September through October.

All three reduced models provided a fit as good as that of our full model (Table 5); therefore we chose the model that had reduced classes for season and area because it was the simplest. Four contrasts of interest were examined:

1 Test for seasonal differences in the odds of observing a set with high bycatch given the set was sampled east of the Mississippi River;

2 Test for seasonal differences in the odds of observing a set with high bycatch given the set was sampled west of the Mississippi River;
3 Test for area differences in the odds of observing a set with high bycatch given the set was sampled in the early season; and
4 Test for area differences in the odds of observing a set with high bycatch given the set was sampled in the late season.

The contrasts were written similarly to those for the "full model." For the four contrasts, we used an α level of 0.025 as significant.

Of the four contrasts, two were significant. The first indicated that in areas east of the river, the odds of observing a set with high bycatch was significantly different between sets sampled in the early season and sets sampled in the late season (Wald χ^2=11.41, df=1, $P>\chi^2$= 0.0007). The odds of observing a set with high bycatch east of the river in the early season was 11 times greater than the odds of observing a set with high bycatch in the late season. The second significant contrast indicated that for sets sampled in the early season, the odds of observing a set with high bycatch east of the river was significantly different from the odds of observing a set with high bycatch west of the river (Wald χ^2=7.99, df=1, $P>\chi^2$=0.0047). In the early season, the odds of observing a set with high bycatch east of the river was 2.7 times greater than observing a set with high bycatch west of the river during the same period.

Table 5

Comparison of reduced logit models with the full model. Diff = difference between full model and reduced model.

	Likelihood ratio statistics			Model comparisons		
	df	G^2	$P > G^2$	Diff (df)	Diff (G^2)	*P*-Value
Full model Season= Apr–Jun, Jul–Aug, Sep–Oct Area = 11–12, 13–14, 15–16, 17–18	12	17.82	0.1213			
Reduced model I Season=spring, summer, fall Area = 11–12, 13–18	6	9.76	0.1352	6	8.06	0.237
Reduced model II Season=Apr–Aug, Sep–Oct Area = 11–12, 13–14, 15–16, 17–18	8	9.99	0.2659	4	7.83	0.098
Reduced model III Season= Apr–Aug, Sep–Oct Area = 11–12, 13–18	4	7.54	0.1100	8	10.28	0.246

Bycatch species associations

Of the 62 species groups observed, 20 occurred in two or fewer sets. The most frequently occurring species were Atlantic cutlassfish, *Trichiurus lepturus* (44% of sets), Atlantic croaker, *Micropogonias undulatus* (38% of sets), Spanish mackerel, *Scomberomorus maculatus* (36% of sets), sand seatrout, *Cynoscion arenarius* (35% of sets), and gafftopsail catfish, *Bagre marinus* (34% of sets). In terms of total abundance (Table 6), Atlantic croaker, sand seatrout, and Atlantic bumper, *Chloroscombrus chrysurus,* accounted for 71% of the total releasable bycatch.

Species included in the main table were Atlantic croaker, sand seatrout, crevalle jack, *Caranx hippos*, gafftopsail catfish, Spanish mackerel, and Atlantic cutlassfish (Table 6). Species included as supplementary variables were striped mullet *Mugil cephalus*, unidentified requiem sharks, gulf butterfish, *Peprilus burti*, cownose ray, *Rhinoptera bonasus*, spotted seatrout, *Cynoscion nebulosus*, Atlantic bumper, blacktip shark, *Carcharhinus limbatus*, red drum, *Sciaenops ocellatus*, unidentified penaeid shrimp, hardhead catfish, *Arius felis*, brown shrimp, *Panaeus aztecus*, cabbage head jellyfish, *Stomolophus meleagris*, bull shark, *Carcharhinus leucas,* and unidentified tonguefish (Soleidae) (Table 6).

The fate of releasable bycatch Correspondence analysis on the fate-by-species table for the entire fishing season indicated that the first two axes explained 97% of the total inertia (conceptually similar to variance) and offered a good representation of the fate-species associations. From the two-dimensional plot (Fig. 6) we discerned three major and one minor groupings.

Species primarily associated with being released dead or disoriented were unidentified requiem sharks, red drum, crevalle jack, and bull sharks. Species secondarily associated with being released dead or disoriented were cownose rays and blacktip sharks. These last two species were primarily associated with being released healthy and appeared to form their own minor grouping.

Table 6

Species used in correspondence analyses. (M) signifies species used in main table and (S) in supplementary table (see "Materials and methods" section. Areas (11–12, 13–14, etc. are shown under each season. Counts are number of organisms observed. Unid. = unidentified.

	Spring				Summer				Fall				
	11–12	13–14	15–16	17–18	11–12	13–14	15–16	17–18	11–12	13–14	15–16	17–18	Total
Atlantic croaker (M)	132	20	50	1	554	1,604	1,732	212	40	97	51	612	5,105
sand seatrout (M)	104	26	17	0	813	1,500	572	2	91	91	197	53	3,466
gafftopsail catfish (M)	255	70	72	24	161	0	96	48	8	28	33	1	796
Atlantic cutlassfish (M)	41	28	209	22	5	1	53	55	7	0	13	36	470
crevalle jack (M)	71	12	133	17	5	0	31	9	22	41	8	0	349
Spanish mackerel (M)	22	11	34	33	29	0	56	33	0	7	6	10	241
Atlantic bumper (S)	0	0	0	0	0	0	2,166	330	0	0	1	0	2,497
striped mullet (S)	344	31	0	0	511	0	0	0	0	9	0	0	895
red drum (S)	21	0	34	0	23	0	24	6	0	12	9	116	245
hardhead catfish (S)	36	5	3	0	100	3	38	3	12	1	3	2	206
tonguefish spp. (S)	0	0	2	0	0	0	1	0	0	0	200	1	204
blacktip shark (S)	37	0	54	1	20	0	19	52	0	0	0	1	184
cownose ray (S)	27	2	26	0	3	0	4	7	0	0	1	0	70
brown shrimp (S)	0	0	1	2	0	0	39	0	0	0	10	13	65
cabbbagehead jellyfish (S)	0	0	5	12	0	0	33	0	1	0	4	6	61
unid. requiem sharks (S)	0	0	23	2	3	0	16	0	0	6	1	6	57
spotted seatrout (S)	2	8	0	0	31	0	0	0	0	0	0	0	41
gulf butterfish (S)	2	0	3	0	0	0	6	3	0	0	0	26	40
bull shark (S)	5	0	0	0	31	0	2	0	0	0	0	1	39
unid. penaeid shrimp (S)	0	0	1	9	5	0	0	0	1	0	0	16	32
Others (not used in CA)	80	5	86	11	147	5	56	24	13	8	72	9	516
Column Total	1,179	218	753	134	2,441	3,113	4,944	784	195	300	609	909	15,579

The second group, species primarily associated with being gilled, were Atlantic croaker, sand seatrout, and unidentified tonguefish. Other species that were associated with being gilled were unidentified penaeid shrimp, Atlantic cutlassfish, gulf butterfish, and Atlantic bumper. These four species were also associated with the third group, species kept by the crew, and those whose fate was unknown. Other species associated with this third group were hardhead catfish, brown shrimp, Spanish mackerel, gafftopsail catfish, striped mullet, and the cabbage head jellyfish.

Temporal and spatial patterns of bycatch species

Spring Correspondence analysis of area by species for spring indicated that the first two axes explained 97% of the inertia and offered a good representation of species-area associations. From the two-dimensional plot (Fig. 7), we discerned three major groupings. The first axis separates the eastern areas of the fishery (zone groups 11–12 and 13–14) from the western areas (zone groups 15–16 and 17–18). The second axis also separates zone group 15–16 from zone group 17–18.

The eastern areas are associated with Atlantic croaker, gafftopsail catfish, sand seatrout, hardhead catfish, bull shark, striped mullet, and spotted seatrout.

Zone group 15–16 is primarily associated with unidentified tonguefish, red drum, crevalle jacks, blacktip shark, unidentified requiem sharks, gulf butterfish, and Atlantic cutlassfish. Zone group 17–18 was associated with Spanish mackerel, brown shrimp, cabbage head jellyfish, and unidentified shrimp.

Summer For the summer, correspondence analysis indicated that the two axes accounted for 94% of total inertia. As in spring, three major species area groupings were observed (Fig. 8). Notable differences in these grouping were that zone group 15–16 appeared to be closer to the eastern groups (13–14 and 11–12). Furthermore, group 13–14 was separated further from zone group 11–12.

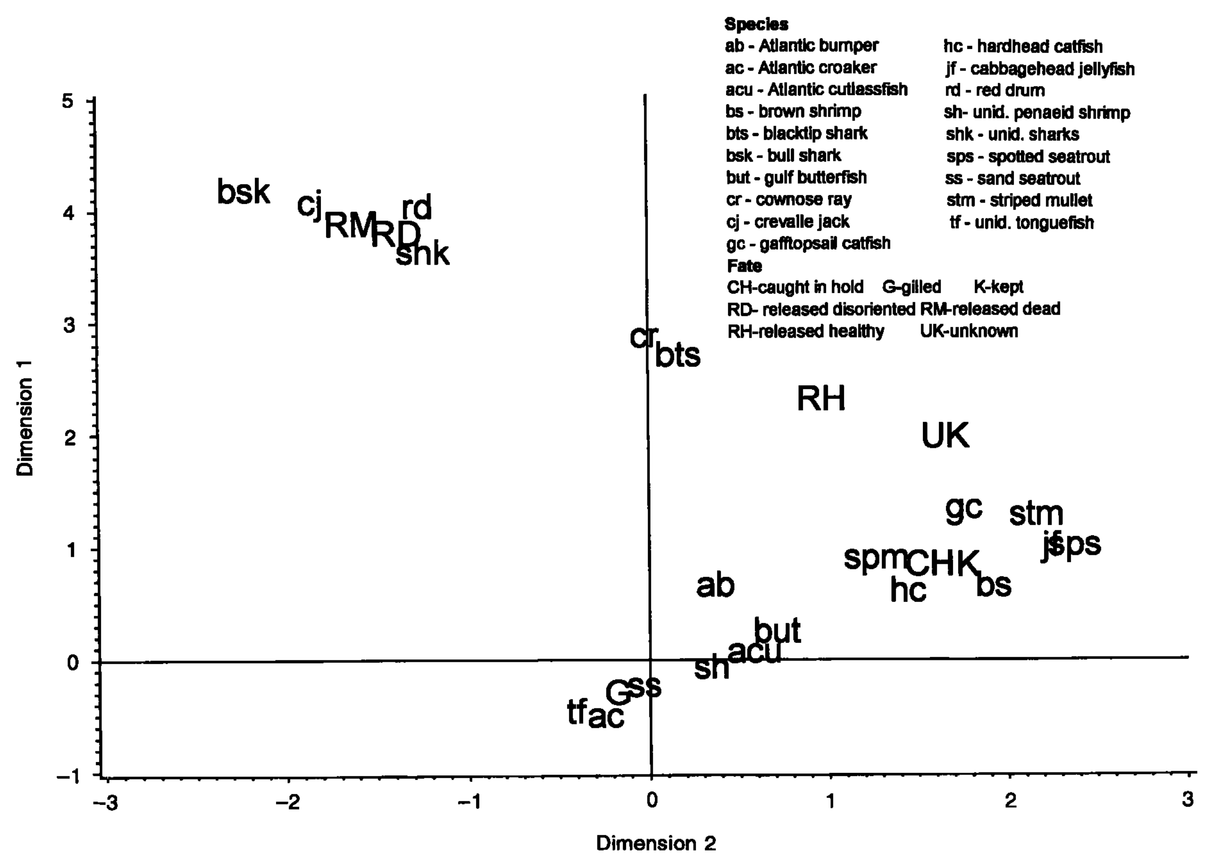

Figure 6

Correspondence analysis plot for species-fate associations for 1995 fishing season.

Zone groups 11–12 and 13–14 were primarily associated with sand seatrout and, to a lesser extent, Atlantic croaker. In addition, zone group 11–12 was also associated with bull shark, striped mullet, spotted seatrout, unidentified shrimp, and hardhead catfish.

Zone group 15–16 was also associated with Atlantic croaker. Other species associated with this area were cabbage head jellyfish, brown shrimp, unidentified tonguefish, Atlantic bumper, unidentified requiem sharks, gulf butterfish, and crevalle jack. Zone group 17–18 was associated with Spanish mackerel, Atlantic cutlassfish, cownose ray, blacktip shark, red drum, and gafftopsail catfish. Secondarily associated species with this area were gulf butterfish and crevalle jack.

Fall The first two axes of the fall analysis explained 97% of the total inertia, once again presenting a good fit of the table. By fall, two major groupings were observed (Fig. 9). Areas from zone group 11–12 to 15–16 formed one group and were separated from the most westerly zone group 17–18.

Zone group 17–18 was primarily associated with Atlantic croaker, red drum, bull shark, blacktip shark, gulf butterfish, unidentified shrimp, and, to a lesser extent, cabbage head jellyfish, brown shrimp, spotted seatrout, Spanish mackerel, and unidentified requiem sharks.

Zone groups 11–12 to 15–16 were associated with hardhead catfish, gafftopsail catfish, sand seatrout, and, to a lesser degree, Spanish mackerel. In addition zone group 13–14 was also associated with crevalle jack and striped mullet and zone group 15–16 with cownose ray, unidentified tonguefish, and Atlantic bumper.

Discussion

Bycatch studies of the menhaden industry were conducted in 1894 on the Atlantic menhaden, *Brevoortia tyrannus*, fishery (Christmas et al., 1960) and in 1948 in the U.S. Gulf of Mexico (Miles and Simmons, 1950). Automatically retained bycatch percentage estimates

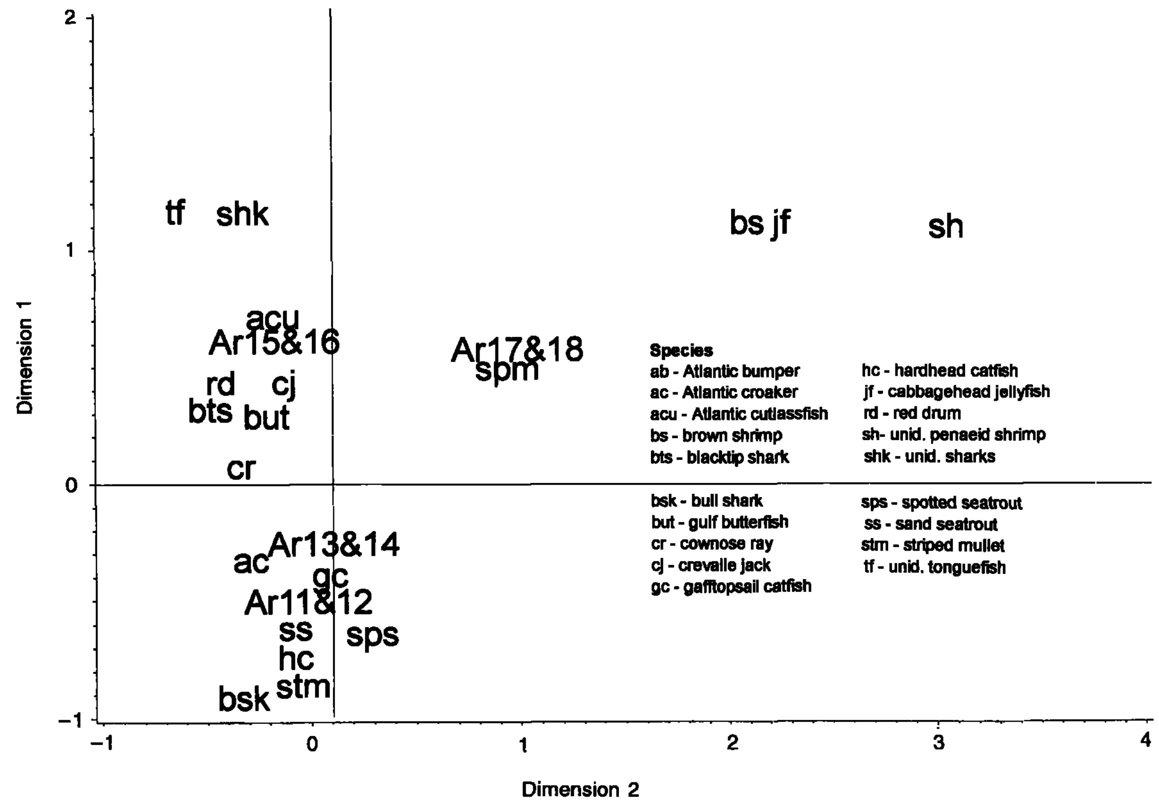

Figure 7

Correspondence analysis plot for species-area associations (spring) for 1995 fishing season.

have ranged from 0.05% (Dunham, 1972) to 3.90% (Christmas et al., 1960) by number, and 1.0% (Condrey[1]) to 2.80% (Christmas et al., 1960) by weight. However, these values are based on bycatch retained in the fish hold. No estimates of the releasable bycatch are available. Based on our analysis, releasable bycatch estimates for the U.S. Gulf of Mexico menhaden fishery range from 0.033% (median) to 0.17% (mean) and reflect the strong positively skewed distribution of the bycatch. Values based on the winsorized mean are intermediate to those of the mean and median and are associated with a lower standard deviation than that for the mean.

As a result of the patchy distribution of menhaden bycatch, examination of the relationship of bycatch to other factors is made more complex. Even after the transformation of our data, gross violations of the ANOVA model assumptions made it an unsatisfactory technique. Because we could not find a suitable transformation, our solution to examining such data would be to convert the variable of interest into a categorical variable and to use categorical techniques in analyzing the data. In our case, the use of loglinear models to identify statistically important interactions was found to be a useful tool in exploring such data. This solution can be considered to fall between studies that can use ANOVA techniques (e.g. Andrew et al., 1995) and those based on the modified negative binomial model as used by Perkins and Edwards (1996).

Legendre (1987) noted that the responses of living organisms to environmental change is nonlinear and in instances nonmonotonic. As loglinear models are insensitive to the shape of the relationship among the variables, Legendre (1987) noted that they are well suited for examining nonmonotically related variables. A further advantage of this type of analysis is that because biological variables respond to interacting environmental variables, they can be used to examine such relationships in detail. By using loglinear models we can include a set of potential interactions in our saturated model, and through a stepwise selection procedure, find those interactions that are statistically important. In our study, bycatch was the issue of interest and the variable that we treated as a response. In effect, we were trying to

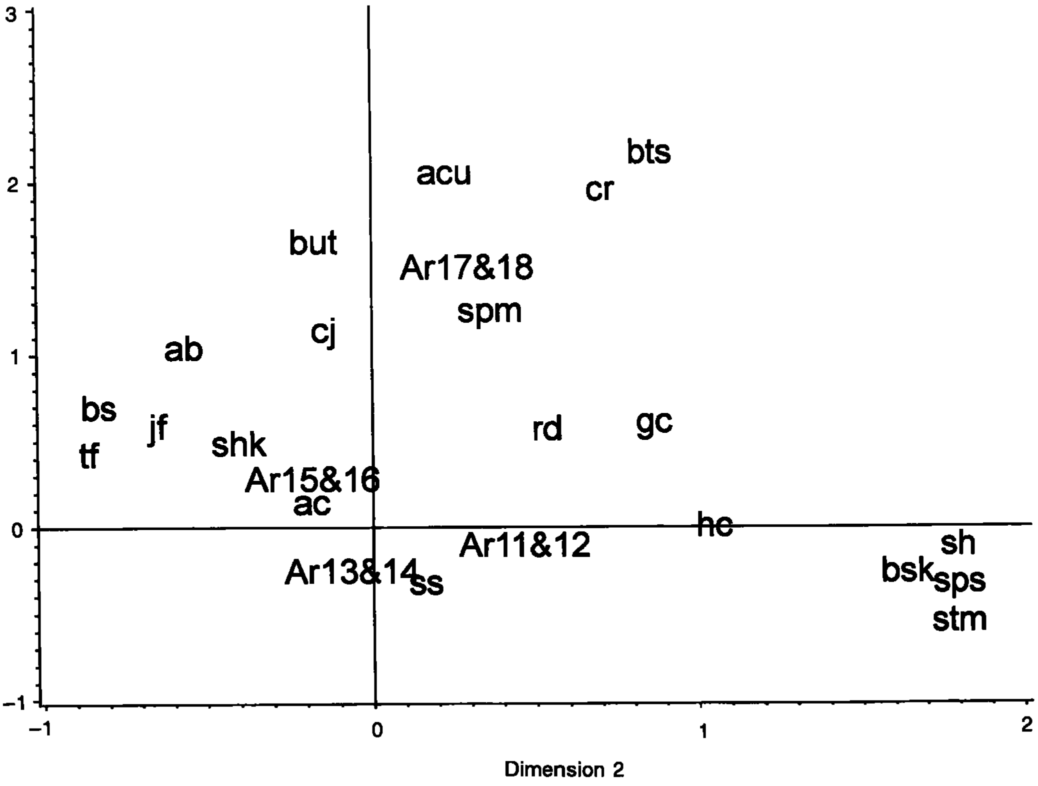

Figure 8

Correspondence analysis plot for species-area associations (summer) for 1995 fishing season.

find factors that could explain the bycatch, and we used loglinear models to find a suitable model that was associated with occurrences of bycatch greater than the median level for the fishery.

By using both loglinear models and logit models with categorical explanatory variables, we have used the loglinear model for model selection and the logit model for a detailed examination of the model of our choice in a manner analogous to ANOVA.

With the stepwise selection procedure, [SAB SAD] was found to be the most suitable loglinear model. We had hypothesized that the presence of dolphins in the fishing area could indicate the presence of high bycatch. Had this been true, the presence of dolphins in the vicinity could have been used by fishermen to avoid setting nets in certain areas. However, such a relation between dolphins and bycatch was found not to exist.

At first glance the issue of bycatch in Gulf of Mexico menhaden fishery may seem to be negligible, given the low bycatch percentage. However the fishery had the second highest annual U.S. commercial landings of 472,000 t in 1995 (U.S. Dep. Commerce, 1996). Further, given the strong positively skewed distribution of releasable bycatch, a small percentage of the total fishing effort would account for much of the take. By structuring our analysis around bycatch rates that were greater than the median, we have attempted to identify the potential "hot spots" in the fishery in terms of areas and seasons. One of the solutions for reducing bycatch in the fishery is to identify such areas and thereby offer the industry a tool for managing their take of bycatch by minimizing fishing effort in these "hot spots."

Our philosophy has been that consideration of all bycatch as a single entity is not the best approach. Although we felt the need to address the total bycatch in our first analysis, we also wanted to take individual species into consideration. We approached this multispecies aspect of our study using correspondence analysis, which we used to identify the commonly associated species in different zones and areas, as well as the different fates. Our results suggest that this approach can have general appeal not only in identifying areas and species of concern but also in suggesting approaches to solutions. For example, our "hot spot" fishing zone and season were

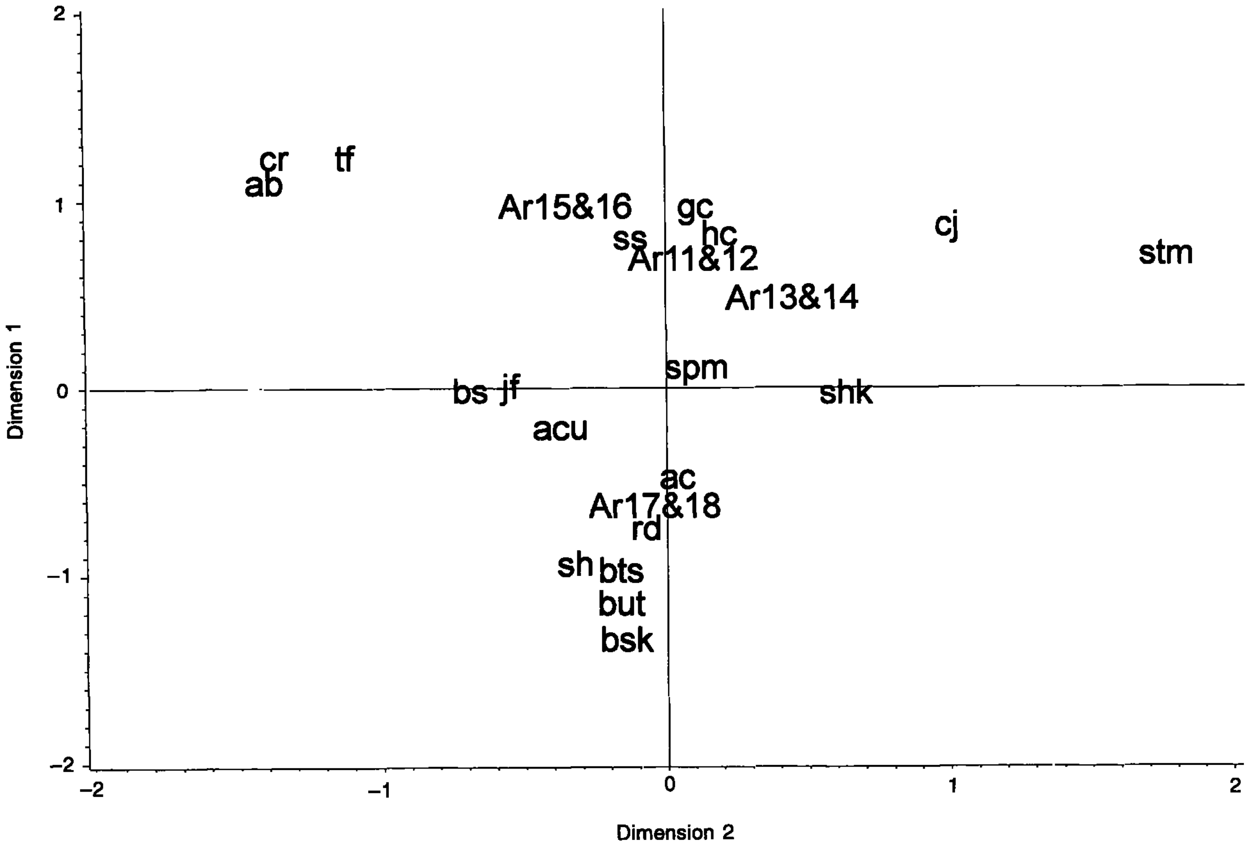

Figure 9

Correspondence analysis plot for species-area associations (fall) for 1995 fishing season.

east of the Mississippi River and April–August. This hot spot was associated with the bycatch of Atlantic croaker, sand seatrout, hardhead catfish, spotted seatrout, and bull sharks. Of these, Atlantic croaker and sand seatrout were the most commonly occurring species in the releasable bycatch and were associated with being gilled. If a reduction in the mortality of these species in the menhaden bycatch were necessary, our study suggests it would require gear modification in the purse seine. A species associated with our hot spot, more likely to require attention, is the bull shark given its life history characteristics. Species-fate associations indicated that bull sharks were primarily associated with being released dead. If a reduction in the mortality of bull sharks as menhaden bycatch were mandated, solutions should be centered around reducing the number of sharks released dead. Rester (1996) has suggested that this could be achieved through modifications to the fish pumping equipment for fish 1 meter in length or larger.

Acknowledgments

We wish to thank the personnel of Gulf of Mexico menhaden industry and members of the Menhaden Advisory Committee, Gulf States Marine Fisheries Commission, for assistance and advice. We wish to thank Donna Rogers, Coastal Fisheries Institute, Louisiana State University, for suggestions and review of the manuscript. We are indebted to James Geaghan, Brian Marx, and Raul Macchiavelli, Department of Experimental Statistics, Louisiana State University, for statistical advice and Steven Branstetter for his review of the manuscript. The manuscript greatly benefitted from the comments of an unknown reviewer and S.X. Cadrin.

Data used in these analyses were collected during a study on menhaden bycatch funded by the National Marine Fisheries Service, MARFIN program (Award No. NA47FF0020).

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