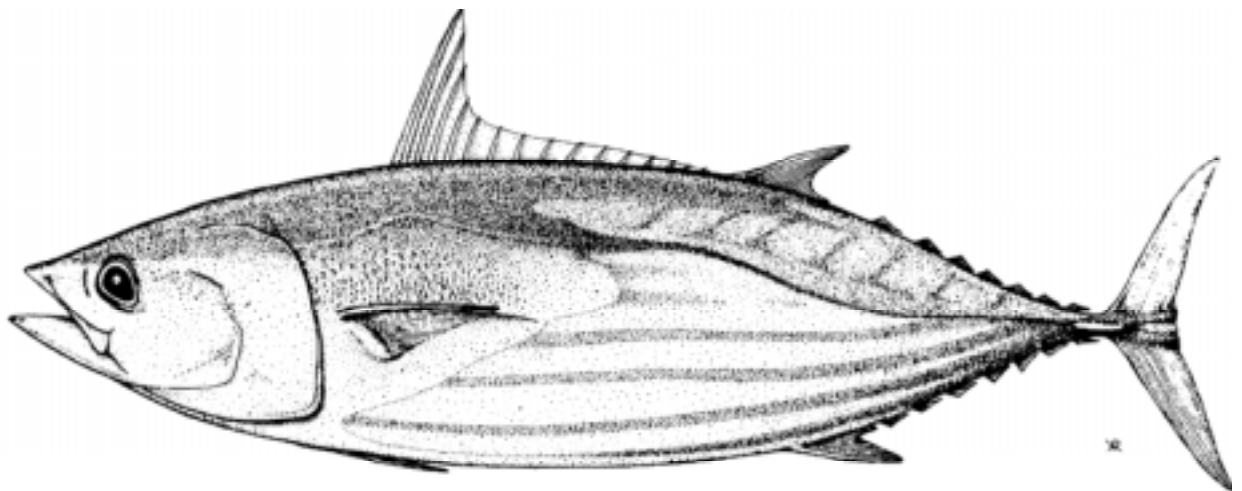




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METHODS FOR ANALYSING BYCATCHES WITH OBSERVER DATA



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INTRODUCTION

The SPC Oceanic Fisheries Programme (OFP) holds observer data covering tuna fishery bycatches in the western and central Pacific Ocean (WCPO) (Figure 1) that have been provided by regional and national observer programmes (Lawson 2001). These data cover longliners (1987–2000), pole-and-line vessels (1998) and purse seiners (1994–2000). Coverage of the catch of target species in the WCPO during the periods covered by the observer data have been 0.18 percent for longline, 0.20 percent for pole-and-line and 3.90 percent for purse seine.

Three of the objectives for which the observer data were compiled are (a) to determine factors that affect catch rates of non-target species; (b) to estimate annual catches of non-target species in the WCPO and the uncertainty of annual catch estimates; and (c) to examine the relationship between coverage rates and the uncertainty of annual catch estimates. This document reviews the literature on methods for conducting these three types of analysis. The studies primarily concern tuna fisheries, but studies on other fisheries have been included when the methodology or results may be applicable. Abstracts of the papers are presented below, in chronological order, followed by summaries for each of the three types of analysis.

ABSTRACTS OF STUDIES OF BYCATCHES BASED ON OBSERVER DATA

Perkins & Edwards (1996)

Perkins & Edwards (1996) examined discards of tuna in the purse-seine fishery in the eastern tropical Pacific Ocean using a method that may also be applicable to the estimation of catches of non-target species. Data covering the discards of tuna per set were collected by observers from September 1989 to March 1992. Observer coverage during the first eleven months was about half of the sets, while coverage during the remaining twenty months was complete. Observer estimates of discards were rounded to integer values, with the rounding interval increasing with the amount discarded.

Discards per set were modelled with the negative binomial with added zeros (NBAZ) distribution, a mixture of a negative binomial distribution and a discrete probability mass at zero, which was appropriate given the large number of zero values. The NBAZ distribution has three parameters (a mixing parameter p and the mean μ and variance α of the negative binomial). Initial fits were made for each of three set types (dolphin sets, log sets, school sets) with no dependence on geographic area, then dependence on three geographic areas was added progressively using stepwise likelihood-ratio tests. Standard errors were computed using (a) a large-sample normal approximation and (b) a bootstrap procedure. Pooled estimates of discard per set were calculated as the weighted average of area-specific estimates, where the weight were proportional to total effort in each area.

With no dependence on area, the maximum likelihood estimator (MLE) of the mean discard per set, for each set type, is simply the sample mean. With complete dependence on area, the MLE for each area reduces to the sample mean in that area and the pooled estimate is computed as a weighted average. In both cases, the variance of the MLE can be estimated using the sample variance. When only the mixing parameter depends on area, the MLE for mean discard per set in each area depends on the ratio of the number of positive observations to the total number of observation in each area. The estimate of the mean discard per set, $(1-p)\mu$, can be more precise than the model parameters.

Geographic area was found to be significant for log and school sets, but not dolphin sets. The mixing parameter for dolphin sets was precise, reflecting the high number of zero observations,

while the negative binomial parameters were imprecise, reflecting the small number of positive observations. In contrast, for log sets, the mixing parameter converged to zero, reflecting the large number of positive sets, which effectively collapsed the NBAZ to a negative binomial. The negative binomial parameters were precise, with coefficients of variation less than 8.5 percent. The results for school sets were intermediate. Mean discard per set for log sets, school sets and dolphin sets was estimated to be 10.5 tons, 1.16 tons and 0.06 tons respectively.

The usual models for integer-valued data, such as the Poisson distribution, did not fit the data because of extreme skewness. The NBAZ is more flexible and is applicable to any set of integer-valued data that exhibit a large proportion of zero observations combined with long positive tails. With the NBAZ, it is possible to examine whether differences in mean discard are attributable to different proportions of zero observations or to different distributions of positive observations.

Lawson (1997)

Annual catches in purse-seine and longline fisheries in the SPC Statistical Area (Figure 2) were estimated by Lawson (1997). Annual purse-seine catches were estimated by multiplying the mean catch rate (tonnes per set), determined from observer data, by the total annual fishing effort (number of sets). The observer data covered 1,377 sets collected during 1994–1996 from Japanese, Korean, Taiwanese and United States purse seiners. Annual catch estimates were derived for 24 species and species groups, including target species (skipjack, yellowfin and bigeye). Catch rates and the total annual number of sets were determined for combinations of fleet and two types of school association (associated and unassociated). The observer data for 1994–1996 were insufficient to estimate annual catch rates for each year, so the data for 1994–1996 were pooled. The total number of sets for each combination of fleet and school association were estimated from the total catch of target species by applying the proportion of the catch by school association and the average catch per set, both of which were determined from logsheet data covering 1996. The annual catch estimates were therefore approximations reflecting catch rates during the previous three years and fishing effort in the previous year.

Catches of non-target species were estimated to have accounted for 0.71 percent of the total purse-seine catch, including discards. The percentage for associated schools was 0.91 percent and for unassociated schools was 0.50 percent. Standard errors of the catch estimates were determined from standard errors of the estimated mean catch rates, which in turn were determined from the sample variances. The coefficient of variation of the catch of non-target species by all four fleets combined was 13.5 percent; however, the coefficients of variation were variable among individual species or species groups.

Annual longline catches were estimated by taking multiplying the mean catch rate (number of fish per 100 hooks), determined from observer data, by the total annual effort (hundreds of hooks). The observer data covered 739 sets collected during 1992–1996 from longliners of China, Fiji, Japan, New Caledonia, Tonga, Taiwan and the United States. Annual catch estimates were derived for 52 species and species groups, including target species (albacore, bigeye and yellowfin). The observer data for 1992–1996 were insufficient to estimate annual catch rates for each year, so the data were pooled. The total annual number of hooks for each fleet was estimated by dividing the total catch of target species during 1996 by the average catch rate (kilograms per 100 hooks). The catch rates were determined from observer data for all fleets, except for the Japanese and Taiwanese fleets, for which the catch rates were determined from catch and effort data, stratified by time-area, that were provided by those fishing nations. The annual catch estimates were therefore approximations reflecting catch rates during the previous five years and fishing effort in the previous year.

Catches of non-target species were estimated to have accounted for 42.5 percent of the total longline catch, including discards. The catch of sharks represented 23.0 percent of the total catch. Standard errors of the catch estimates were determined from standard errors of the estimated catch rates, which in turn were determined from the sample variances. The coefficient of variation of the catch of non-target species by all fleets combined was 2.3 percent. The coefficients of variation were variable among individual species or species groups, but generally lower than for estimates of purse-seine bycatches.

Romanov (1997)

Romanov (1997) examined the average bycatch per set and the bycatch per 1000 tonnes of tuna caught by Soviet/Russian purse seiners in the western Indian Ocean using observer data collected from 1986 to 1992. A total of 494 sets were observed, of which 377 sets (76 percent) were positive. The data were insufficient to estimate bycatches on an annual basis, so the data for 1986–1992 were pooled. The data were stratified into four seasons (winter, spring, summer and autumn) and three types of school association (free-swimming schools, whale associated and log associated). The observed catch of tuna (7,252 tonnes) represents 30.2 percent of the total catch of tuna (24,018 tonnes) by Soviet/Russian vessels during 1986–1992 and 0.47 percent of the total catch of tuna (1,540,309 tonnes) by all fishing nations during 1986–1992. The geographic and temporal distribution of the observed catch was representative of the catch by Soviet/Russian vessels, except for the autumn season, during which the number of observed sets was proportionately less than the total. Sets resulting in a catch of less than 0.5 tonne were ignored. The condition of discards, and hence their survival rate, was not considered. Confidence intervals for the average bycatch per set and the bycatch per 1000 tonnes of target species were calculated; the derivation of the confidence intervals was not explained and are assumed to be plus or minus twice the sample standard error. A total of 43 species or species groups were observed.

Bycatch (all species other than skipjack and yellowfin) from free-swimming schools was 3.4 tonnes per 1000 tonnes of target species, or 0.34 percent of the total catch, and consisted primarily of sharks, rays and billfish. Bycatch from schools associated with whales was 11.0 tonnes per 1000 tonnes of target species, or 1.09 percent of the total catch, and consisted primarily of sharks. Bycatch from schools associated with logs was 41.3 tonnes per 1000 tonnes of target species, or 3.97 percent of the total catch, and consisted primarily of rainbow runner, mahi mahi, triggerfish, sharks, wahoo, billfish and mackerel scad. One unidentified marine turtle was observed in a log-associated school. One whale mortality due to entanglement in the net was observed. The bycatch from all school types combined was 27.2 tonnes per 1000 tonnes of target species or 2.65 percent of the total catch. The confidence interval for the bycatch per 1000 tonnes of target species from all school types combined represented 65.4 percent of the point estimate.

Dorn et al. (1997)

Dorn et al. (1997) examined the relationship between the proportion of vessels observed in a fleet and the percent error (upper and lower bounds of 95 percent confidence intervals as a percentage of the point estimate) of bycatch estimates by resampling complete observer data for Alaskan groundfish fisheries. Low levels of observer coverage are associated with relatively high levels of error for most target and non-target species. For target species and some non-target species, error decreases rapidly as coverage increases from 10 percent to 30 percent and much more slowly with further increases of coverage. For some bycatch species, however, uncertainty remains high, even when all vessels are observed.

Vølstad et al. (1997)

Vølstad et al. (1997) examined the relationships between the coefficient of variation for bycatch estimates and (a) the proportion of vessels observed in a fleet and (b) the proportion of hauls sampled, by resampling complete observer data for trawl fisheries for walleye pollock and yellowfin sole in the Bering Sea and Aleutian Islands region. Simulations indicated that variability is high when either the fraction of vessels sampled within the fleet or the fraction of hauls sampled within a vessel is low. With random selection of observed vessels, estimates with acceptable error bounds can be made for frequently occurring species by sampling 30 percent of the vessels. For less frequently occurring species, a much larger proportion of the fleet would need to be sampled. They also examined vessel-specific coefficients of variation and found that they declined rapidly as coverage of the number of hauls sampled per vessel increased up to 30 percent, although the coefficients of variation for a given level of sampling varied considerably among vessels. Catch estimates based on the delta-distribution (Pennington, 1996) were examined for rare species with highly skewed catches per haul.

Turnock & Karp (1997)

Turnock & Karp (1997) conducted simulations to examine the relationship between the proportion of hauls sampled and the coefficient of variation of vessel-specific estimates of salmon bycatch in the trawl fishery for walleye pollock in the Bering Sea and Aleutian Islands region. Within haul sampling variability is a concern when estimating the bycatch of rarely occurring species that are likely to be highly aggregated. Since information on the within haul distribution was unknown, they assumed a Poisson distribution. The results suggest that the coefficient of variation for estimates of the mean number of salmon per haul decreases rapidly as the proportion of the haul sampled approaches 20–30 percent and that variability remains high until 50–70 percent of the hauls are sampled. The distribution of salmon within a haul is almost certainly more complex than was assumed; hence, within-haul variability is likely to be much greater than predicted.

Kleiber (1998)

Kleiber (1998) estimated annual takes and kills of sea turtles in the Hawaiian longline fisheries using a statistical model to predict takes from logbook data. The statistical model was parameterised using observer data covering 1,922 longline sets made during 1994–1996. Variables that were considered for the model included geographic and temporal location (latitude, longitude, year, month, time of day); vessel and gear attributes (vessel length, target species, number of hooks per set, number of hooks between floats, bait species, number of light sticks); and environmental attributes (sea surface temperature, sun elevation, moon phase). The catches of 13 other species was also considered. Variables were tested for significance with classification trees. Total kills were estimated by multiplying the estimated number of takes by an estimate of the probability of a kill given a take and the condition of the turtle recorded by the observer (hook ingested; hooked, but hook not ingested; hook in unknown location; alive and uninjured at release; dead; condition not recorded).

Uncertainty in the point estimates of takes and kills was assessed using the distribution obtained from 1,000 bootstrap estimates, where the bootstrap samples were drawn by trip, rather than by set. The calculation of the kill per take was included in the bootstrap procedure.

For loggerheads, sea surface temperature was the most important explanatory variable, followed by the catch of albacore. When sea surface temperature was excluded because it was not well-represented in the logbook data, latitude was the primary explanatory variable. For olive ridleys, the

catch of yellowfin was the most important variable, but the relationship was not statistically significant, so no explanatory variables were included. For leatherbacks and green turtles, no significant explanatory variables were included.

The prediction intervals of kills in 1997 were between 102 and 284 percent of the point estimates: i.e. the 95 percent prediction interval for loggerhead kills in 1997 was 23–74, compared to the point estimate of 50; 0.1–20 leatherback kills compared to the point estimate of 7; 10–47 olive ridley kills compared to the point estimate of 28; and 0.03–0.94 green turtle kills compared to the point estimate of 0.5.

Hay et al. (1999)

Hay et al. (1999) estimated bycatches of eulachons (*Thaleichthys pacificus*), an anadromous smelt, in the trawl fishery for shrimp in British Columbia using two types of ratios of the eulachon catch to the shrimp catch that were determined from observer data. Total bycatches were estimated as the eulachon to shrimp catch ratio multiplied by the total shrimp catch. For the estimate based on the 'MI-ratio' (mean in-season catch ratio), the catch ratio was calculated as the mean of the catch ratios for observed tows. For the estimate based on the 'PI-ratio' (pooled in-season catch ratio), the observer data were pooled and the catch ratio was calculated as the catch of eulachons from all observed tows divided by the catch of shrimp from all observed tows. The two in-season estimates based on catch ratios were compared to estimates based on the product of catch rates determined from observer data and total effort determined from logbooks that become available at the end of the season. It was initially believed that the MI-ratio estimate, based on individual tows, would be preferable since it would allow measures of uncertainty of the catch ratio to be determined from the replicates. However, the MI-ratio estimates appeared to inflate bycatch estimates, while the PI-ratio estimates were in closer agreement with the estimates based on catch rates. This may be due to the fact that short tows that have high catches of eulachons are given the same weight in the MI-ratio as longer tows with little or no bycatch. Since fishing effort is concentrated in areas with high catch rates of target species and low catch rates of non-target species, the MI-ratio may not be representative of the fishery.

Francis et al. (1999, 2000)

Francis et al. (1999) estimated annual catches on non-target fish species in the New Zealand longline fishery from the 1988–89 fishing year (October 1 to September 30) to the 1996–97 fishing year by multiplying mean catch rates (number of fish per 1000 hooks), determined from observer data, by the total effort (thousands of hooks). The analysis was updated in Francis et al. (2000) for 1988–89 to 1997–98. The observer data cover 2,604 sets made during 99 trips from 1986–87 to 1997–98. The coverage rate was 7.5 percent of the total number of hooks set, although annual coverage averaged 4.2 percent from 1988–89 to 1991–92 and then 23.4 percent from 1992–93 to 1997–98. Observer coverage during the 1986–87 and 1987–88 fishing years was 0.2 and 0.9 percent respectively and these levels were considered too low to be representative; hence, these fishing years were not considered. The number of hooks sets in the New Zealand EEZ declined by 88 percent from the early 1980s to the late 1990s, as foreign vessels left the fishery.

The data were stratified into two fleets, (i) foreign and chartered vessels and (ii) domestic vessels, and two geographic areas, (i) north and south of 38°00'S on the west coast and (ii) north and south of 43°50'S on the east coast. The foreign and chartered vessels fished only a short season (April to July) and while domestic vessels fished a longer season (mainly from January to June), observer coverage of domestic vessels was mostly limited to the period April–June; therefore, the data were not stratified by season. Seventy species of fish were observed. The mean catch rate and total

numbers caught were estimated for the 17 most common species. Estimates of the 95 percent confidence intervals were derived from 1000 bootstrap samples. Estimates of catch rates were plotted by fishing year to determine whether there were trends over the ten-year period or differences among strata. Estimates of the logarithm of catch rates were plotted against latitude.

For blue shark, porbeagle shark and mako sharks caught during the 1996–97 and 1997–98 fishing years, the total catch in weight was estimated from estimates of the total numbers caught using length-weight regressions to convert length frequencies to weight frequencies. The catch in weight was estimated separately for each of the two regions and, for blue shark and porbeagle shark, for each sex. A similar procedure was used for Ray's bream, except only 1997–98 was considered and the length frequency for the southern region was applied to both regions.

Discriminant function analysis was used to determine whether longline sets that caught striped marlin could be distinguished from those that did not. Eleven variables related to community structure (Shannon-Weaver diversity index, number of species caught); the environment (sea surface temperature, standard deviation of sea surface temperature along the longline, moon phase); gear attributes (soak time, maximum hook depth, percentage of hooks baited with squid, percent baited with fish); and the total number of fish caught. There were 91 marlin sets and 140 non-marlin sets used in the analysis, which supported the hypothesis that sets catching striped marlin can be distinguished from those that do not. The error rate for classifying marlin sets was 14 percent, whereas the error rate for classifying non-marlin sets was 39 percent.

From 1994–95 to 1997–98, the species composition changed as foreign and chartered vessels left the fishery and domestic vessels entered. For most species there were large differences in catch rates between the domestic fleet and the foreign/chartered fleet and/or between the northern and southern regions. Long-term trends in catch rates were noted for only a few species in some strata. Inter-annual variation in catch rates was high.

For five species of tuna and broadbill swordfish, confidence limits for the estimates of numbers of fish caught in all strata combined during 1997–98 ranged from 23.9 (southern bluefin tuna) to 56.1 (bigeye tuna) percent of the point estimates. For five shark species, the confidence interval ranged from 26.1 (porbeagle shark) to 131.6 (school shark) percent of the point estimate. For six species of other fish, the confidence interval ranged from 32.8 (Ray's bream) to 95.3 (oilfish) percent of the point estimate.

McCracken (2000)

McCracken (2000) extended the work of Kleiber (1998) in estimating annual takes of sea turtles in the Hawaiian longline fisheries, using a statistical model to predict takes from logbook data. The statistical model was parameterised using observer data collected during 279 trips taken from 1994 to 1999. The annual observer coverage was less than 5 percent of trips. Variables that were well-represented in the logbook data, and hence were considered for the model, included geographic and temporal location (latitude, longitude, distance to 17°C isotherm, distance to 19°C isotherm, year, month, day); vessel and gear attributes (vessel length, trip type, number of hooks per set, number of hooks between floats); and environmental attributes (sea surface temperature). Certain variables were pooled into categories or included as polynomials. The catches of 12 other species was also considered. Classification trees were used to explore variables associated with turtle take and to determine possible categorical splits for continuous variables and pooling of categories for categorical variables; the response variable was defined as 'take' or 'no take'. The final models were developed with generalised linear models (GLMs) and generalised additive models (GAMs). All quantitative variables were scaled to eliminate problems associated with different units of

measurement. For GLMs and GAMs, quasi-likelihood estimation was used in a log-linear Poisson model of takes. Smoothing splines were used in the GAMs. The Bayesian information criterion (BIC) was used for stepwise selection.

The independence of observations of the take per set for green turtles, leatherbacks and olive ridleys was assumed on the basis of the small number of trips during which more than one take occurred. Although 29 trips out of 55 with positive loggerhead takes had takes in more than one set, loggerheads were modelled at the set level. Prediction intervals were approximated using a bootstrapping algorithm, although, for loggerheads, the error structure of the data was modelled, instead of assuming a Poisson distribution.

Total kills were estimated by multiplying the estimated number of takes by an estimate of the probability of a kill given a take, which was determined from the observer data.

The species of ten unidentified turtles were assigned on the basis of the sea surface temperature and the latitude for the observation. Two takes of unidentified turtles for which sea surface temperature and latitude were not clear indicators of species were split fractionally between three species.

The explanatory variables included in the prediction models were as follows. For loggerheads, month in three categories, latitude as a polynomial and sea surface temperature in two categories were included; for olive ridleys, sea surface temperature in two categories; for leatherbacks, latitude in four categories; for green turtles, no explanatory variables were included.

The prediction intervals of kills in 1999 were between 116 and 300 percent of the point estimates: i.e. the 95 percent prediction interval for loggerhead kills in 1999 was 28–102, compared to the point estimate of 64; 1–27 leatherback kills compared to the point estimate of 11; 11–96 olive ridley kills compared to the point estimate of 55; and 1–19 green turtle kills compared to the point estimate of 6.

Baird & Bradford (2000)

Baird & Bradford (2000) examined factors that may have influenced seabird bycatch taken by longliners in New Zealand waters using (a) exploratory analysis, (b) generalised linear models (GLM), and (c) paired significance tests.

Exploratory analysis consisted primarily of the examination of the degree of overlap of approximate confidence intervals (point estimate plus or minus two standard errors) of mean catch rates determined from observer data that were stratified, for several factors, into categories. The factors were examined separately for Japanese vessels and New Zealand vessels. The factors of geographic or temporal location that were examined included fishing year (1986–87 to 1997–98 for Japanese vessels, 1991–92 to 1997–98 for New Zealand vessels); month; geographic area (six areas); and 1° latitude band (29°S to 48°S). Vessel and gear attributes included vessel (30 Japanese vessels, 17 New Zealand vessels); vessel length; vessel gross registered tonnage; engine power; vessel category (New Zealand vessels only: four categories); mainline type (Japanese vessels only: kuralon, nylon multifilament, mix, unknown); line setting speed; use of tori poles (yes, no, unknown); tori line length (11 categories for Japanese vessels, six categories for New Zealand vessels); number of streamers (21 categories for Japanese vessels and 13 categories for New Zealand vessels); bait type (four categories of the percentage of squid per set); use of bait casting machines (yes, no); number of snoods per basket; average snood length; set start time; time of capture of dead bird; set duration; and haul duration. Environmental attributes included moon phase and sea surface temperature (22

categories). Geographic area, time of capture and moon phase were examined by year. The use of tori poles was examined by geographic area.

A GLM with a negative binomial error distribution was fitted to the number of birds caught per set. The negative binomial was used to overcome the over-dispersion noted when using Poisson models for similar data. A negative binomial fitted the overall distribution well, but when fitting the GLM, the parameters did not converge, probably because the bird count distributions are extreme cases of the distribution, with a high fraction of sets with no bird captures and occasional extremely large values. Also, the count distributions did not appear to have the same shape in different areas and years. A binomial model was therefore used to model sets that caught or did not catch birds. The Poisson error distribution was used since over-dispersion decreased as the amount of variance explained increased. A multinomial model (multiple logistic regression model), in which bird captures can be described as (0, 1, ..., many), was tested. However, each number allowed in the distribution added another set of parameters and the Akaike Information Criterion (AIC) rapidly increased. For fitting models, the data were grouped into three time periods and three areas. Binomial and Poisson GLMs of increasing complexity were applied to the bird capture data. These models were generally ill-fitting, explaining a small percentage of the original deviance, and the Poisson models were over-dispersed. Both models suggested where significant differences might lie, but the significance level was dubious.

Paired significance tests were conducted with (a) the large sample test for difference in success probabilities for binomial data and (b) a two-sided randomisation test for differences in mean number of birds per set and mean number of birds per 100 hooks. Examples of variables tested include time period, area, whether tori lines were used, whether sets were made at night, moon phase, sea surface temperature, and interaction terms.

Effects found to influence bird catch rates for observed Japanese vessels included area, time period, moon phase and sea surface temperature. The use of tori lines and setting at night did not show up as statistically significant. The domestic vessel data were all in the northern area and year group 1993–94 to 1997–98 and birds were caught during hauling as well as setting. No significant effects due to moon phase, the use of tori lines, and setting at night were found when all bird encounters were included. When only dead birds were considered, they appeared more likely to be caught around the time of the full moon.

Lennert-Cody (2000)

Lennert-Cody (2000) examined the effects of sample size on estimates of bycatches in the tuna purse-seine fishery in the Eastern Pacific Ocean using resampling of observer data. The observer data covered sets on floating objects from 1995 to 1998, during which annual coverage was greater than 90 percent. Observer coverage levels of 5, 10, 15, 20, 25, 30, 40 and 50 percent were simulated by resampling trips, without replacement. For each level of observer coverage, 1000 samples of trips were generated. Estimates of the total bycatch were calculated as (a) the product of the bycatch per set (the sum of bycatch in a stratum divided by the number of sets) and the total number of floating object sets in the fishery and (b) the product of the bycatch per ton of tuna caught (the sum of bycatch in a stratum divided by the tonnage of tuna caught) and the total tonnage of tuna caught in association with floating objects in the fishery. The data were stratified into six species groups (billfishes, sharks, rays, dolphinfish, wahoo, other large fish) and four years (1995–1998).

Simulation results were summarised with (a) box and whisker plots of the estimates of total bycatch, (b) the ratio of the mean squared error of total bycatch estimates to the known bycatch, which approximates the coefficient of variation when bias is small, and (c) estimates of the statistical

power (the probability that the null hypothesis of no difference will be rejected when it is false) to distinguish between two estimates of total bycatch. Differences in total bycatch for a species group and level of coverage were simulated by multiplying estimates of total bycatch by a scale factor that ranged from 1.1 to 2.0, in intervals of ten percent.

Sampling coverage of less than ten percent tended to result in distributions of total bycatch estimates that were highly skewed towards large estimates.

The approximate coefficient of variation decreased rapidly with increasing sampling coverage up to 20 and 25 percent. For coverage levels of 25 and 30 percent, the approximate coefficient of variation was 10 percent for billfish, 20–30 percent for sharks, rays and dolphinfish, and 40 percent for other large fish.

The power to distinguish between estimates of total bycatch for a 40 percent increase in the catch of sharks and billfishes, with observer coverage of 30 percent and an alpha of 0.1, was greater than 0.75. For a 60 percent increase in dolphinfish and rays, coverage of 30 percent resulted in power of greater than 0.80. For a 100 percent increase in other large fish, 30 percent coverage resulted in power of less than 0.80. For some species groups, the results varied considerably among years.

Large values of observed bycatch, some the result of converting tons to numbers of fish, tend to result in large estimates of total bycatch at low levels of coverage and thus greater variance and lower estimates of power. Additional reductions in variance might be achieved by stratification of the data by geographic area and time period.

SUMMARY

Factors Affecting Bycatch

Observer data usually include information on geographic and temporal location, vessel and gear attributes, and environmental attributes. The manner in which the studies described above have examined the relationship between bycatches and various factors can be summarised as follows:

- Perkins & Edwards (1996) used stepwise likelihood ratio tests with the NBAZ distribution to examine observer data covering the purse-seine tuna fishery in the eastern Pacific Ocean and found that geographic area was significant for log and school sets, but not dolphin sets.
- Lawson (1997) stratified observer data covering longliners and purse seiners in the WCPO tuna fishery by fishing nation and, for purse seiners, by school association.
- Romanov (1997) stratified observer data covering purse seiners in the Indian ocean tuna fishery into four seasons and three school associations.
- Kleiber (1998) used classification trees to examine the effect of several variables on turtle takes in the Hawaiian longline tuna fishery. For loggerheads, sea surface temperature was the most important explanatory variable, followed by the catch of albacore. When sea surface temperature was excluded because it was not well-represented in the logbook data, latitude was the primary explanatory variable. For olive ridleys, the catch of yellowfin was the most important variable, but the relationship was not statistically significant, so no explanatory variables were included. For leatherbacks and green turtles, no significant explanatory variables were included. Bycatch estimates were determined for calendar years. McCracken (2000) extended Kleiber's analysis and found that, for loggerheads, month in three categories, latitude as a polynomial and sea

surface temperature in two categories were important as explanatory variables. For olive ridleys, sea surface temperature in two categories, and, for leatherbacks, latitude in four categories were important, whereas for green turtles, no explanatory variables were important.

- Francis et al. (1999, 2000) stratified observer data from the New Zealand tuna longline fishery into two fleets, two geographic areas and fishing year.
- Baird & Bradford (2000) examined factors that may have influenced seabird bycatch taken by longliners in New Zealand waters using (a) exploratory analysis (primarily the degree of overlap of approximate confidence intervals for mean catch rates), (b) generalised linear models (GLM), and (c) paired significance tests. Effects found to influence bird catch rates for observed Japanese vessels included area, time period, moon phase and sea surface temperature. The use of tori lines and setting at night did not show up as statistically significant. For New Zealand vessels, no significant effects due to moon phase, the use of tori lines, and setting at night were found when all bird encounters were included. When only dead birds were considered, they appeared more likely to be caught around the time of the full moon.

Estimation of Annual Bycatches from Observer Data

Types of estimates of bycatches

Bycatches have been estimated from observer data by (a) multiplying catch rates determined from observer data by estimates of total fishing effort (Lawson, 1997; Romanov, 1997; Francis et al. 1999, 2000); (b) multiplying catch ratios determined from observer data by the total catch of target species (Hay et al., 1999; see also Vølstad et al., 1997, and Lennert-Cody, 2000); or (c) predicting the catch per set with a model parameterised with observer data and applied to logbook data (Kleiber, 1998; McCracken, 2000).

Lack of sufficient data for annual estimates

The lack of sufficient data to determine estimates for each year of the period covered by the observer data was a problem encountered in several studies. When this occurred, the data were either pooled over the entire period (Lawson, 1997; Romanov, 1997) or the years for which the data were insufficient were ignored (Francis et al. 1999, 2000).

Large number of zero observations

Observations of catches of non-target species are characterised by a large number of observations of zero catch per fishing operation (longline set or purse-seine set). For bycatches measured in units of numbers of individuals, a common model for integer-valued data, the Poisson distribution, does not fit the data well because of extreme skewness. When considering discards per set, which also exhibit a large number of zero observations, Perkins & Edwards (1996) considered three distributions with a probability mass at zero and concluded that the negative binomial with added zeros (NBAZ) was the most appropriate. For estimates of bycatches, Vølstad et al. (1997) examined the delta distribution, which is a lognormal distribution with a spike at zero.

Evaluating uncertainty

The degree of uncertainty of bycatch estimates has been evaluated with standard errors or confidence intervals based on a large-sample approximation (Lawson, 1997; Romanov, 1997) or a bootstrap procedure (Francis et al. 1999, 2000; Kleiber, 1998; McCracken, 2000). The lack of the assumed relationship between the mean and variance implies that large-sample approximations may

not be appropriate to determine second-order properties of parameter estimates. McCracken (2000) considered a Poisson distribution for turtle take per set and found overdispersion, i.e. the variance was larger than, rather than equal to, the mean.

Vølstad et al. (1997) examined standard errors based on two-stage cluster sampling (Cochran, 1977), with and without the delta distribution, for standard two-stage estimators and ratio estimates. The application of the delta distribution in the standard two-stage estimators resulted in the highest coefficients of variation for estimates of the catch by species. The lowest coefficients of variation resulted from the ratio estimator with the delta distribution.

Independence of observer data

Large-sample approximations may also not be appropriate if there is a lack of independence; however, the independence of the observed catches per set for non-target species in tuna fisheries has not been examined directly. McCracken (2000) assumed the independence of observations of the take per set for green turtles, leatherbacks and olive ridleys on the basis of the small number of trips during which more than one take occurred.

Selecting models

The lack of the assumed relationship between the mean and variance or the lack of independence also implies that large sample approximations may not be appropriate for selecting models. McCracken (2000) used stepwise selection based on a generalised information criterion (GIC), but only as an aid in determining the final predictive model, since, inter alia, assuming sets within trips are independent is questionable.

Mean catch rates vs pooled catch rates

Catch rates estimated as the mean of the observed catches per set have an associated standard error determined from the observed variance in the catch rates among sets, whereas catch rates estimated by pooling, i.e. by dividing the sum of the catch from all observed sets by the number of the observed sets or hooks, do not. Hay et al. (1999) found the different methods for estimating catch rates resulted in large differences and suspected that the mean catch rate was biased, whereas the pooled catch rate was not. In any case, resampling procedures, such as the bootstrap or jackknife, can be used to estimate standard errors when pooling.

Use of explanatory variables

For the Hawaii longline fishery, Kleiber (1998) and McCracken (2000) examined several explanatory variables, but only those that were well-represented in logbook data, since the model was developed to predict turtle takes for sets covered by logbook data, but not observer data. However, for many fisheries, including most tuna fisheries, the coverage by logsheet data is low and the logsheet data do not include many explanatory variables other than temporal and geographic location. In these cases, bycatches are estimated by multiplying observed catch rates by the total effort or observed catch ratios by the total catch of target species (Lawson, 1997; Romanov, 1997; Hay et al. 1999; Francis et al. 1999, 2000). Since the statistics on total effort or catch are usually stratified by only a small number of variables (i.e. fishing nation, temporal and geographic location, and, for purse seine, school association, and, for longline, set depth as indicated by the number of hooks between floats), the observed catch rates are stratified by a similar small number of variables, even though the observer data may contain a much greater amount of information on vessel and gear attributes and environmental conditions.

Observer Coverage Rates and the Uncertainty of Bycatch Estimates

Stages of the sampling design

Coverage rates for the sampling of bycatches by observers can be considered for several stages of the sampling design. For the purse-seine tuna fisheries in the WCPO, for example, the stages could include (1) the fishing nations, (2) the vessels within each fishing nation, (3) the trips taken by each vessel, (4) the sets during each trip, (5) the brails per set and (6) the fish per brail. Vølstad et al. (1997) examined the relationships between the coefficient of variation for bycatch estimates and (a) the proportion of vessels observed in a fleet (which corresponds to stage 2 above) and (b) the proportion of hauls sampled (stage 4) for trawl fisheries. Lennert-Cody (2000) examined the relationship between three measures of uncertainty of bycatch estimates and the percentage of trips sampled (i.e. stages 1–3 above were considered as a single stage) for purse-seine tuna fisheries in the eastern Pacific.

Relationship between the observer coverage rates and the uncertainty of bycatch estimates

Simulations conducted by Vølstad et al. (1997) indicated that with random selection of observed trawlers, estimates with acceptable error bounds can be made for frequently occurring species by sampling 30 percent of the vessels. For less frequently occurring species, a much larger proportion of the fleet would need to be sampled.

Vølstad et al. (1997) examined vessel-specific coefficients of variation and found that they declined rapidly as coverage of the number of hauls sampled per vessel increased up to 30 percent, although the coefficients of variation for a given level of sampling varied considerably among vessels.

Simulations of observer coverage of purse-seine sets on schools associated with floating objects in the eastern Pacific tuna fishery conducted by Lennert-Cody (2000) indicated that coverage of less than ten percent tended to result in distributions of total bycatch estimates for species groups that were highly skewed towards large estimates. The approximate coefficient of variation decreased rapidly with increasing sampling coverage up to 20 and 25 percent. For coverage levels of 25 and 30 percent, the approximate coefficient of variation of the catch estimates was 10 percent for billfish, 20–30 percent for sharks, rays and dolphinfish, and 40 percent for other large fish.

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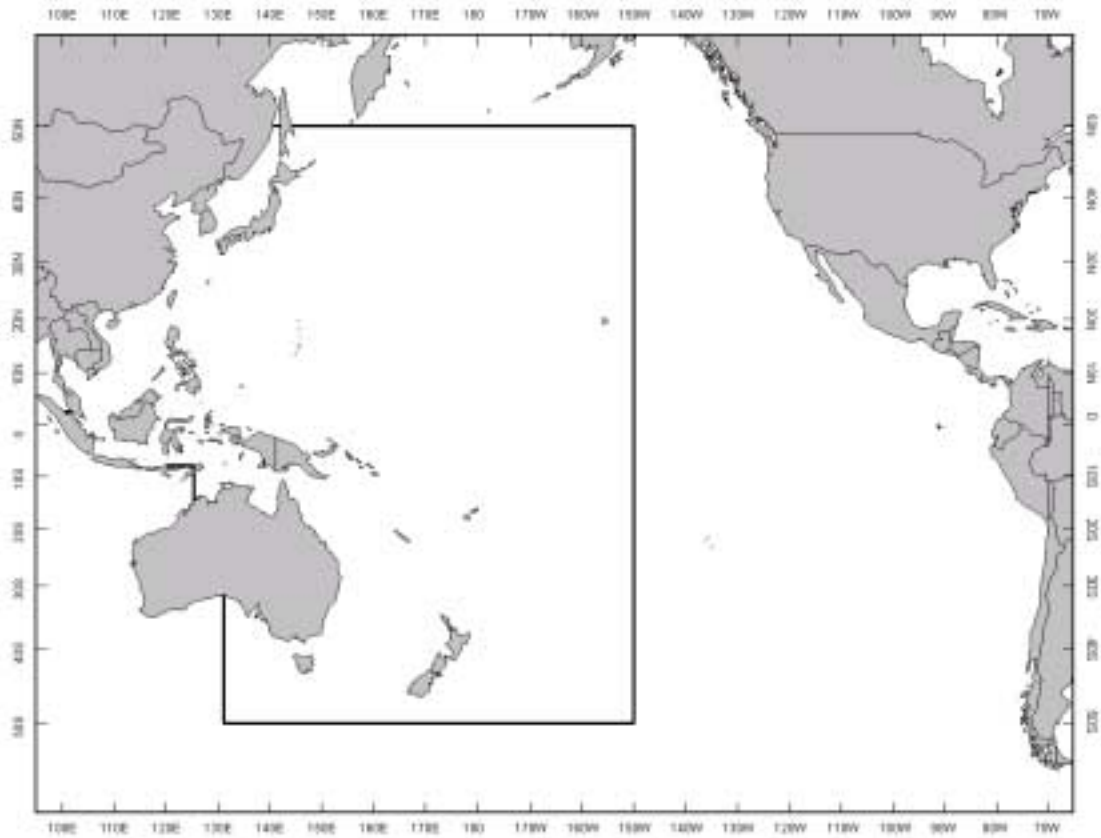


Figure 1. Western and Central Pacific Ocean (WCPO) Area

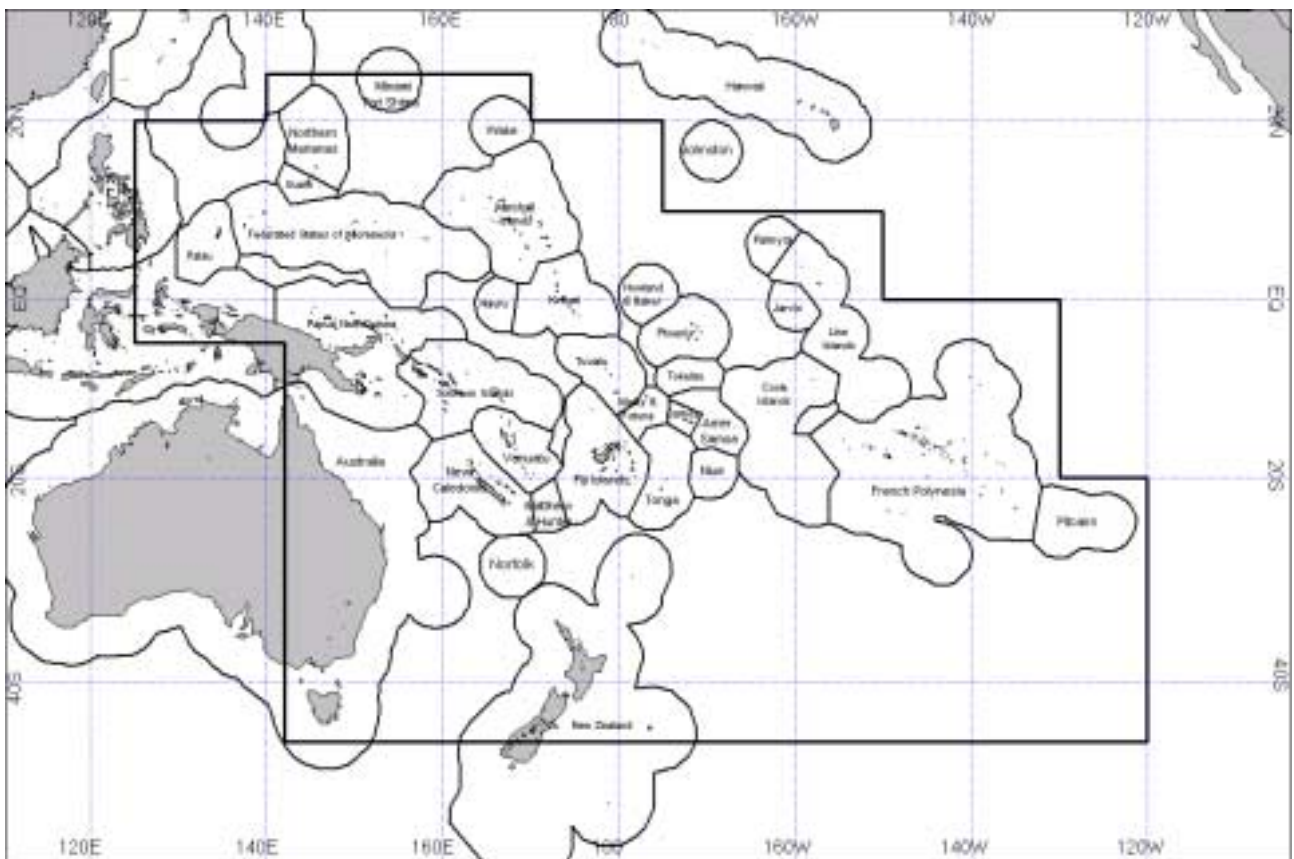


Figure 2. SPC Statistical Area