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Quantifying tag reporting rates for Atlantic tuna fleets using coincidental tag returns

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Abstract – Uncertainty about reporting rates of tags returned by fishermen has often prevented tagging data from being used in stock assessments. In this study we conduct a meta-analysis to estimate tag reporting rates of commercial tuna fleets by comparing their tag return data with those of the USA longline pelagic observer program. The longline fleets of Venezuela and the USA are estimated to report about 0.8% of tags caught, compared with less than 0.1% for Canadian, Spanish and Japanese longline fleets. For some fleets with sparse return data or for those that do not overlap often with the observer fleet, reporting rate estimates are sensitive to changes in the spatio-temporal resolution over which comparisons are made. Regardless of these sensitivities, the estimated reporting rates are low and there are likely to be large differences in reporting rate between different combinations of flag and gear.

Key words: Fisheries management / Stock assessment / Mark-recapture / Tag recovery / Tuna / Billfish / Sharks

Résumé – L'incertitude liée aux taux de retour des marques rapportées par les pêcheurs a souvent empêché d'utiliser les données de marquage pour les évaluations de stocks. Dans cette étude, nous effectuons une méta-analyse pour estimer les taux de retour des marques des flottilles commerciales thonières, en comparant les taux de retour de marques avec ceux de programme d'observation de la pêche palangrière des Etats-Unis. Les flottilles de palangriers du Venezuela et des USA rapportent environ 0,8 % des marques capturées, comparées à moins de 0,1 % pour les flottilles palangrières canadiennes, espagnoles et japonaises. Pour quelques flottilles dont les données de retour de marques sont sporadiques ou celles qui correspondent peu souvent avec celles de la flottille d'observation, les estimations des taux de retour sont sensibles aux changements de résolution spatio-temporels pour laquelle ces comparaisons sont faites. Sans tenir compte de ces variations, les taux de retour sont faibles et probablement fort différents selon les différentes combinaisons possibles entre pays et engins de capture.

1 Introduction

Tagging data have been used to evaluate hypotheses about behaviour (Block et al. 1998; Sedberry and Loefer 2001), fishing mortality (Hearn et al. 1998; Latour et al. 2003), natural mortality (Hoenig et al. 1998; Gaertner and Hallier 2003) and the temporal and spatial distribution of fish populations (Dupuis 1995; Block et al. 2002). Since 1950 over 500 000 Atlantic sharks, billfish and tunas have been marked with conventional tags. It has proven difficult to use these data in stock assessments other than to provide general information about movement and growth of individuals. Survival rate can be estimated from tagging data without knowledge of reporting rates. However reporting rate allows survival to be divided into two components of central concern in stock assessment: natural mortality and exploitation rate. Due to the paucity of information allowing the separate estimation of tagging parameters, such as tag induced mortality rate and reporting rate, these have remained uncertain, which has often prevented tagging data from being used in stock assessment. The aim of this paper is to quantify reporting rates for Atlantic tuna fleets and, in doing so, allow tagging data to be used more effectively in stock assessment.

In this study, reporting rate is defined as the probability that a tag is observed, reported and recorded in the ICCAT (International Commission for the Conservation of Atlantic Tunas) conventional tagging database, given that it is caught. A wide range of experiments and methods have been devised to estimate reporting rate (Pollock et al. 1995; 2001; 2002b). In theory both time-invariant tag reporting rate and tag retention-survival rate can be estimated solely from tagging data (Brownie et al. 1985), although in practice reporting

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Table 1. The species, flags and gear groups included in this analysis.

Species	Flags	Gear Groups
Albacore (ALB)	Canada (CAN)	Baitboat (BB)
Bigeye tuna (BET)	Spain (ESP)	Longline (LL)
Bluefin tuna (BFT)	France (FRA)	Purse seine (PS)
Blue shark (BSH)	Ghana (GHA)	
Blue marlin (BUM)	Japan (JPN)	
Portbeagle shark (POR)	USA	
Sailfish (SAI)	USA observer (OBS)	
Skipjack tuna (SKJ)	Venezuela (VEN)	
Shortfin mako (SMA)		
Swordfish (SWO)		
White marlin (WHM)		
Yellowfin (YFT)		

rates may vary over time and estimates may have relatively low precision (Hoenig et al. 1998). It follows that additional information regarding reporting rate is required, usually from a reward-based study or observer programme. The multiple-component approach to reporting rate estimation was first applied by Paulik (1961) and then by Kimura (1976). The underlying theory is that the voluntary reporting rate of the commercial fleet may be informed by comparing their tag return rate (tags caught per fish caught according to the information provided to the ICCAT by commercial fisheries) and that of an operator whose reporting rate is 100 percent (i.e. when observers or cooperative fishermen are reporting every tag that they observe; Pollock et al. 2002b). By analysing data collected after harvesting, reporting rate estimation is separated from confounding quantities that determine the probability of recapturing a tag but operate prior to recapture, such as the rates of tag shedding, tag induced mortality, tag failure and natural mortality.

The data available for this study include incomplete records of age of individuals at release and recapture, which preclude the use of more sophisticated approaches that could account for varying fractions of tagged fish among cohorts and age-dependent reporting rates (e.g. Hearn et al. 1999 and Pollock et al. 2002a). In this research we used a relatively simple multi-component extension of Paulik's (1961) method to estimate reporting rates for 18 fleets that fish for sharks, billfish and tunas in the Atlantic (similar to the approach of Heifetz and Maloney 2001). The method makes use of the data of the USA longline pelagic observer program (1990-2000) and approximately half of the conventional return data (1954-2008) for pelagic fish species that ICCAT has recently made publicly available (data of the US pelagic longline observer program are currently unavailable, due to confidentiality issues).

2 Materials and methods

2.1 Theory

In this research, reporting rates are quantified using the theory described by Kimura (1976):

$$\frac{T_{com}}{C_{com}} = m\lambda_{com} \tag{1}$$

where T is the number of tags returned, C is the number of fish caught, m is the mark rate (defined here as the probability of catching a tag given that a fish is caught), λ is the reporting rate and the subscript *com* refers to a commercial fishing fleet.

It follows that mark rates can be obtained from the data of an observer fleet *obs*, whose reporting rate is assumed to be 100 percent, or a commercial fleet whose reporting rate is known:

$$m = \frac{T_{obs}}{C_{obs}} = \frac{T_{com}}{\lambda_{com}C_{com}} \tag{2}$$

and therefore:

$$\lambda_{com} = \frac{T_{com}C_{obs}}{T_{obs}C_{com}}. \tag{3}$$

Using this model, there were four areas of methodological investigation based on the following questions: (1) What is an appropriate spatio-temporal resolution for comparing the return rates of different fleets? (2) How should tag returns be modelled to account for possible non-independence among these observations? (3) How should reporting rates be modelled for different flags, gears and, potentially, species? (4) How can mark rates be modelled to aid the estimation of reporting rates?

The estimation was undertaken in a Bayesian framework to provide probabilistic estimates of reporting rates that may be used as priors in later analyses. Markov Chain Monte Carlo simulation using the Gibbs sampler was undertaken using R 2.92 (R Development Core Team 2006), the "R2WinBUGS" package and WinBUGS 1.43 (Lunn et al. 2000).

2.2 Data analysed

Data from the USA longline observer program, ICCAT conventional tagging database and ICCAT catch and effort database were used to construct a database of tag returns and catch in numbers per species for each flag and gear type, twice a year by 20 × 20 degree ocean square from 1960. Since this date, tags have been returned by over 500 combinations of flag (e.g. USA) and gear (e.g. longline), and attached to over 120 different species of shark, tuna and billfish. For this analysis we included only the top seven flags with the highest contribution, 12 species and 3 gear group codes; this selection encompasses approximately half of the recorded returns (these flags, gear groups and species are listed in Table 1). The rod and reel fisheries were not included in this analysis due to evidence suggesting under-reporting of catches. For example, in many cases, the US rod and reel fleet reported returns of tags where no catch of fish is reported in the ICCAT catch and effort database.

2.3 The spatio-temporal resolution of comparisons

Disaggregating tag return data in space and time reduces the assumption that tags are distributed homogenously. However, disaggregation greatly reduces the number of overlapping tag return events that are used to provide information on reporting rates (see Tables 2 and 3 for a summary of the degree of overlap among flags). Aggregating returns over the entire Atlantic and time series ensures that they all overlap, but may be

Table 2. The number of times that tag return events of different flags occur in the same temporal and spatial unit, given different spatio-temporal resolution of comparisons and minimum days at liberty of tags (aggregated over species and gear groups).

20 × 20 Biannual DAL > 91

	ESP	FRA	GHA	JPN	OBS	USA	VEN
CAN	0	0	0	14	0	7	0
ESP		65	15	67	17	24	5
FRA			13	17	0	1	3
GHA				8	0	6	0
JPN					117	113	28
OBS						157	1
USA							5

40 × 40 Biannual DAL > 91

	ESP	FRA	GHA	JPN	OBS	USA	VEN
CAN	0	0	0	11	0	7	0
ESP		41	19	72	33	46	8
FRA			12	14	0	6	2
GHA				8	0	5	0
JPN					133	126	38
OBS						146	64
USA							47

20 × 20 Biannual DAL > 91

	ESP	FRA	GHA	JPN	OBS	USA	VEN
CAN	0	0	0	12	0	7	0
ESP		44	13	64	16	20	4
FRA			9	15	2	2	1
GHA				8	0	5	0
JPN					119	94	21
OBS						149	8
USA							8

40 × 40 Biannual DAL > 91

	ESP	FRA	GHA	JPN	OBS	USA	VEN
CAN	0	0	0	14	0	8	0
ESP		94	18	83	17	25	6
FRA			21	29	0	2	5
GHA				9	0	7	0
JPN					117	119	30
OBS						158	1
USA							8

Table 3. The spatio-temporal range of tag returns used to provide reporting rates by flag and gear.

Fleet and gear type	Latitude (to 20 × 20 deg square From	To	Latitude (to 20 × 20 deg square From	To	Year From	To	Total coincidental recaptures DAL > 0	DAL > 91	DAL > 182
CAN LL	40	60	–80	–40	1979	1990	3	2	2
CAN PS	–20	60	–100	20	1972	1978	203	120	119
ESP BB	0	60	–40	0	1973	2007	1174	348	303
ESP LL	–60	60	–80	20	1975	2007	28	22	19
ESP PS	–20	60	–80	20	1973	2007	480	117	72
FRA BB	0	60	–40	0	1975	2007	18	3	1
FRA PS	–20	60	–40	40	1975	2007	1571	119	40
GHA BB	–20	20	–20	20	1976	2006	24	20	18
GHA PS	–20	20	–80	20	1976	2006	141	40	11
JPN BB	–40	20	–80	20	1969	1984	17	17	12
JPN LL	–60	80	–100	40	1956	2007	162	145	116
JPN PS	–20	20	–40	20	1967	1991	1	1	1
OBS LL	0	60	–100	–20	1992	2000	3664	2949	2711
USA LL	–40	60	–100	20	1986	2007	287	256	227
USA PS	–20	60	–100	20	1962	2003	386	308	307
VEN LL	–40	80	–100	0	1970	2007	167	137	127
VEN PS	–20	40	–100	20	1982	2007	1	1	1

biased due to regional differences in mark rate. Disaggregating returns at very fine scales makes the criteria for coincidence more stringent, reducing the number of returns that are compared to provide reporting rates. In order to minimise assumptions about the spatio-temporal homogeneity of mark rates, the "base case" comparisons were made at the most detailed level of the database: on 20 × 20 degree ocean squares, biannually (October – March, April – September). The sensitivity of reporting rate estimates to comparisons made at coarser spatio-temporal resolutions is analysed below.

It is likely that tags become better mixed (their distribution is more spatially homogeneous) as time at liberty increases. It is necessary either to deal with the statistical issue of non-independence among short-term return events (due to "clumping" of tags) or to exclude tags captured before a sufficient mixing time. The derivation of defensible models to describe non-independence among return events is problematic since they must account for gear specific phenomena. For example, purse seine fleets targeting tuna aggregations that may have been tagged recently. Such inconsistencies in the historical return data make it difficult to quantify overdispersion which can be incorporated into the negative binomial model to account for non-independence. A second problem with spatially disaggregated comparisons is that reporting-rate estimates by flag and gear can be sensitive to changes in the spatio-temporal resolution of comparisons (Fig. 1). In order to avoid these related problems, the "base case" method applied here simply ignores all returns within the first 91 days at liberty. The sensitivity of reporting rate estimates to the duration at liberty before which returns are ignored is also assessed using sensitivity analysis.

The probability of the observed number of reported tag returns, T, given the mark rate, reporting rate and observed number of fish without tags, C, can be calculated by a negative binomial distribution:

$$P(T|C,\lambda,m) = \prod_t \prod_a \prod_f \prod_g \prod_s \left[\frac{\left(T_{t,a,f,g,s} + C_{t,a,f,g,s} - 1\right)!}{T_{t,a,f,g,s}!\left(C_{t,a,f,g,s} - 1\right)!} \times \left(1 - \left(\lambda_{f,g}\, m_{t,a,s}\right)\right)^{C_{t,a,f,g,s}} \left(\lambda_{f,g}\, m_{t,a,s}\right)^{T_{t,a,f,g,s}} \right] \quad (4)$$

where the subscripts t, a, f, g and s refer to time, area, flag, gear and species, respectively.

The likelihood function of Equation (4) has an equivalent binomial counterpart. During this research we investigated specific cases of the negative binomial distribution that could account for overdispersion, and chose to describe the more general method above. With this approach, the fleets that overlap most with the observer fleet are likely to provide the most reliable information regarding reporting rate (mark rates are assumed to be the same as the observer tag return rate). These fleets then provide mark rates outside of the spatial range of the observer fleet, using extrapolation to obtain reporting rates of those operators that do not coincide with the US longline observer fleet.

2.4 Modelling of mark rate and reporting rate

The mark rate m, defined here as the probability of capturing a tag given that a fish is caught, is estimated for each time, area and species, and prescribed a relatively non-informative beta prior distribution:

$$m_{t,a,s} \sim beta\,(1,1)\,. \quad (5)$$

Various models for reporting rates were considered, including those that model species, flag and gear effects hierarchically (a hyper prior is assigned to each set of effects) and those that include second order interactions (e.g. species × flag). The linear models for reporting rates do not converge without the inclusion flag × gear or flag × species interactions. Convergence diagnostics also indicate that there is insufficient information to support models that include species, flag and gear effects. Similarly, we did not pursue hierarchical models because they do not converge unless strongly informative priors are prescribed for hyperparameters.

It is generally difficult to attribute variation in inferred reporting rate to either gear or species, due to data limitations: in some cases combinations of flag and gear return tags for a limited number of species. The "base case" model estimates reporting rates for each flag and gear (a meta-analysis across species) because subjectively we consider constant reporting rate among species to be a lesser assumption than that of constant reporting rate among gear type groups. This "base case" model choice is supported by the Deviance Information Criterion (DIC): a reporting rate specific to each flag and gear (e.g. USA baitboat, USA longline, Ghanaian baitboat) is the most parsimonious (DIC = 11 461) compared with a model estimating a reporting rates for each flag and species (DIC = 13 556).

Similarly to mark rate, the reporting rate r, defined here as the probability of detecting and reporting a tag given that a tag is caught, is estimated for each flag f, and gear g and given a relatively non-informative beta prior distribution:

$$r_{f,g} \sim beta(1,1). \quad (6)$$

The exception to this is the USA longline observer fleet whose reporting rate is assumed to be 100 percent:

$$r_{f=1,g=1} = 1. \quad (7)$$

2.5 Assumptions and problems

It should be noted that the method applied here relies on a number of important assumptions. (1) Tags are distributed homogenously over the spatial and temporal scale of comparisons among operators. (2) Selectivity over time and among flags and gears remains constant and/or tags are distributed evenly over the length or age classes over which animals are selected. (3) Reporting rates remain constant over time and among species, and are the same in areas beyond spatio-temporal comparisons (e.g. the reporting rate of the US longline fleet is the same in areas where there is no observer coverage). (4) Catches are accurately reported. All of these assumptions are likely to be invalid to some extent. To a variable degree among fleets, reporting rates are obtained on the basis of extrapolation and should consequently be interpreted with caution. The French longline fleet, for example, overlaps relatively infrequently with the American, Japanese or Venezuelan fishing operations, which are directly obtained from the USA longline observer data. It follows that it may be unreasonable to assume the same reporting rates for French longline operations outside of the spatio-temporal range of these overlaps. Table 3 summarises the temporal and spatial range of the return data by fleet.

Hearn et al. (1999) identified the potential biases of aggregating tag returns and catch across age groups (also relevant to aggregation over other subsets of the population). However, the available dataset for Atlantic tunas and sharks is sparse with patchy or no catch-at-age data for most species. Corresponding age data are not available for most of the tag returns either. Were this covariate information available, further disaggregation by age might not be feasible because it would further reduce the already marginal number of comparisons.

Despite these important limitations, we applied this method as no previous work has been done on the estimation of reporting rates of Atlantic fleets and, therefore, a lack of other informative data and methods.

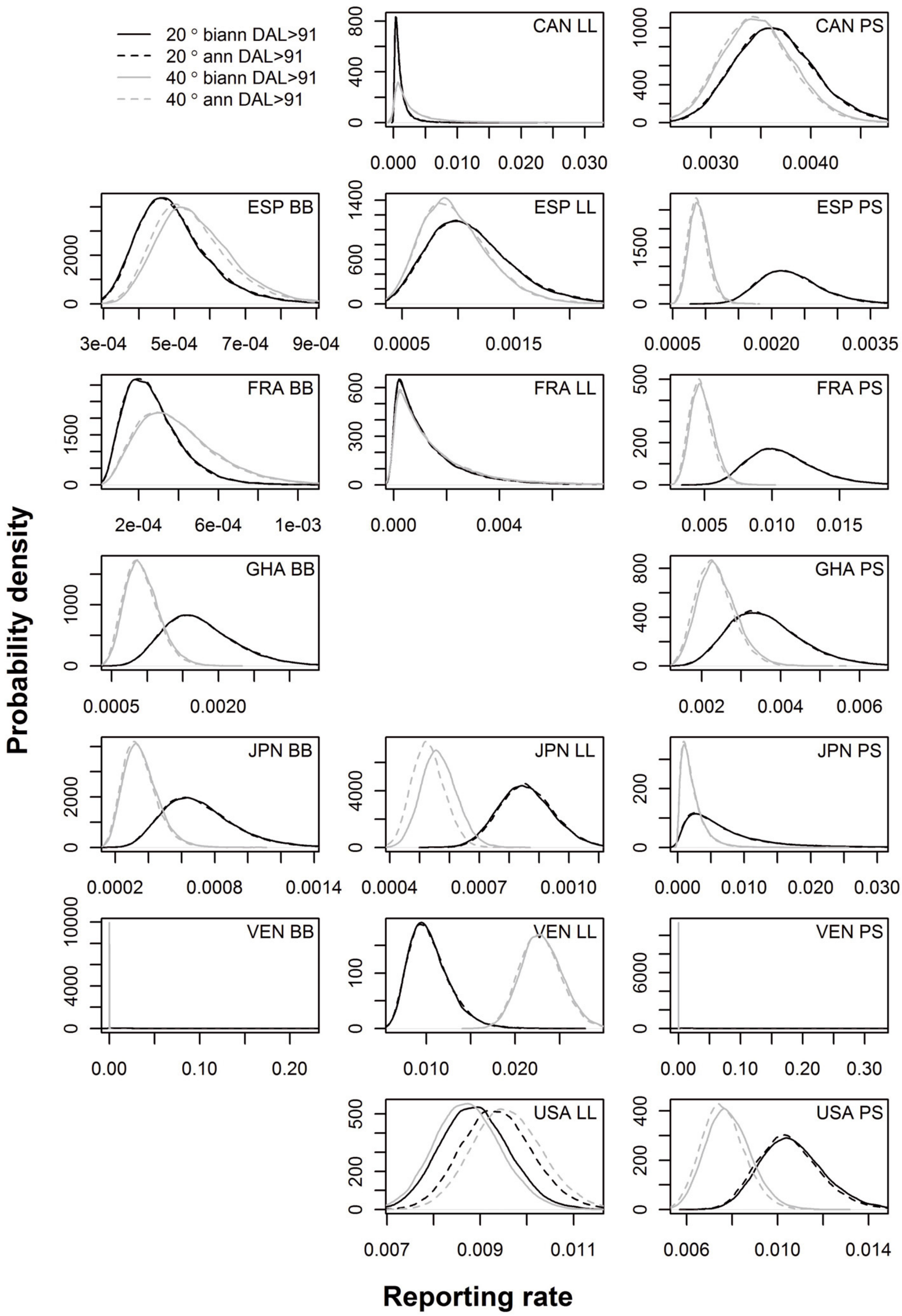

Fig. 1. Marginal posterior density plots of reporting rate comparing the base case model (solid black lines) with plots that model mark rate at different spatio temporal resolution. "20°" and "40°" refer to 20 × 20 and 40 × 40 degree ocean squares, respectively. "biann" and "ann" refer to biannual and annual temporal disaggregation. All tags returned after 91 days at liberty are included ("DAL > 91").

Table 4. Posterior probability distributions of Atlantic reporting rate (percentage of tags 445 reported) under the base case scenario (20 × 20 degree ocean cells, disaggregated biannually) and three alternative runs (CV = standard deviation / mean; DAL = days at liberty).

		20 × 20 biannual DAL > 91					Alternative runs					
		Percentile					DAL > 30		40 × 40		Annual	
Flag	**Gear**	**2.50%**	**50%**	**97.50%**	**Mean**	**CV**	**Mean**	**CV**	**Mean**	**CV**	**Mean**	**CV**
CAN	**LL**	0.013	0.076	0.491	0.120	1.211	0.134	1.225	0.375	1.697	0.120	1.260
CAN	**PS**	0.292	0.363	0.452	0.365	0.111	0.470	0.115	0.348	0.105	0.364	0.110
ESP	**BB**	0.034	0.048	0.075	0.050	0.214	5.533	0.204	0.056	0.195	0.050	0.211
ESP	**LL**	0.052	0.107	0.200	0.112	0.342	0.196	0.292	0.098	0.315	0.113	0.339
ESP	**PS**	0.151	0.225	0.338	0.230	0.209	3.597	0.209	0.091	0.169	0.229	0.207
FRA	**BB**	0.007	0.025	0.062	0.027	0.527	0.110	0.364	0.041	0.520	0.027	0.527
FRA	**LL**	0.003	0.087	0.524	0.132	1.096	10.566	1.082	0.149	1.079	0.133	1.086
FRA	**PS**	0.672	1.045	1.655	1.075	0.233	10.176	0.215	0.485	0.180	1.078	0.234
GHA	**BB**	0.093	0.168	0.297	0.175	0.300	1.119	0.299	0.094	0.260	0.175	0.302
GHA	**PS**	0.211	0.353	0.586	0.365	0.266	3.678	0.240	0.241	0.206	0.365	0.267
JPN	**BB**	0.036	0.068	0.124	0.071	0.314	0.444	0.328	0.036	0.289	0.070	0.311
JPN	**LL**	0.069	0.085	0.105	0.086	0.108	0.104	0.106	0.057	0.104	0.086	0.106
JPN	**PS**	0.063	0.516	2.369	0.693	0.913	9.779	0.909	0.220	0.801	0.682	0.903
VEN	**BB**	0.038	1.734	16.560	3.365	1.379	5.435	0.978	0.008	1.014	3.422	1.381
VEN	**LL**	0.680	1.002	1.584	1.036	0.224	1.372	0.212	2.304	0.103	1.039	0.225
VEN	**PS**	0.156	3.101	24.040	5.340	1.265	1.061	1.371	0.007	0.716	5.349	1.260
USA	**LL**	0.751	0.887	1.039	0.889	0.083	0.921	0.080	0.873	0.082	0.937	0.082
USA	**PS**	0.817	1.056	1.379	1.067	0.135	1.550	0.146	0.784	0.127	1.051	0.133

3 Results

Reporting rate estimates from the base case scenario (20 × 20 degree ocean squares, biannual data) are detailed in Table 4 and illustrated in Figure 2. In general, median reporting rate estimates were less than 4 per cent and excluding the Venezuelan bait boat and purse seine fleets; all remaining reporting rates were less than 1.5%. The six lowest reporting fleets (Canadian longline, Spanish baitboat, French baitboat, French longline, Japanese longline and Japanese baitboat) are estimated to report less than 1 tag in every 1000.

In some cases, spatial disaggregation of the data led to large changes in reporting rate estimates of some operators, in particular French purse seiners, Ghanaian baitboats, Ghanaian purse seiners, Japanese longliners, Senegalese baitboats and Venezuelan purse seiners (Table 4, Fig. 1). These estimates appear sensitive to the assumption of complete tag mixing in the Atlantic after 91 days and should be treated with caution. Due to their greater coincidence with the USA longline observer fleet and more comparable gear selectivities, the reporting rates of the Japanese, Venezuelan and USA longline fleets are likely to be estimated the most reliably (mean estimates are 9, 104 and 89 per 10 000 tags returned, respectively).

Despite the considerable sensitivity of some reporting rate estimates to the level of spatial disaggregation, large differences remain among fleets (Table 4, Figs. 1, 3). For example, regardless of the disaggregation of comparisons, the reporting rate of the USA longline fleet is over four times that of the Spanish longline fleet and their 95% probability intervals do not overlap. The mean reporting rate of the Japanese longline fleet is estimated to be 8–12 times lower than the Venezuelan longline fleet, regardless of the spatio temporal resolution of comparisons or the minimum number of days at liberty of returned tags.

Sensitivity testing indicated that the observer fleet has a greater tendency to catch tags that have been at liberty for longer; this can be observed in Figure 3, where increasing the minimum days at liberty leads to a general tendency for lower inferred reporting rates even in longline fleets whose selectivity is more comparable to that of the observer fleet (a smaller fraction of tags are removed from the observer fleet data as minimum days at liberty is increased).

The estimation of shared mark rates leads to posterior correlation among estimated reporting rates. We include the variance-covariance matrix among logit transformed reporting rates under the "base case" spatially aggregated scenario (Table 5). These posterior distributions are approximated well by the multivariate normal distribution and are provided to enable correct priors to be prescribed where needed in subsequent analyses.

4 Discussion

These analyses represent a preliminary attempt to quantify tag reporting rates for 18 combinations of flag and gear that fish for tunas, billfish and sharks in the Atlantic. These analyses rely on important assumptions. Additionally, reporting rate estimates of some fleets are sensitive to alternative data aggregations and the minimum duration at liberty of tags. However, this research indicates that reporting rates are likely to differ strongly among different combinations of flag and gear. The very low reporting rate of most fleets is of concern since it indicates that opportunistic tagging programs may only be achieving a small fraction of their potential value for stock assessment.

With the exception of two uncertain estimates for the Venezuelan baitboat and purse seine fisheries, median reporting rates were estimated to be below 1% for most fleets, a rate

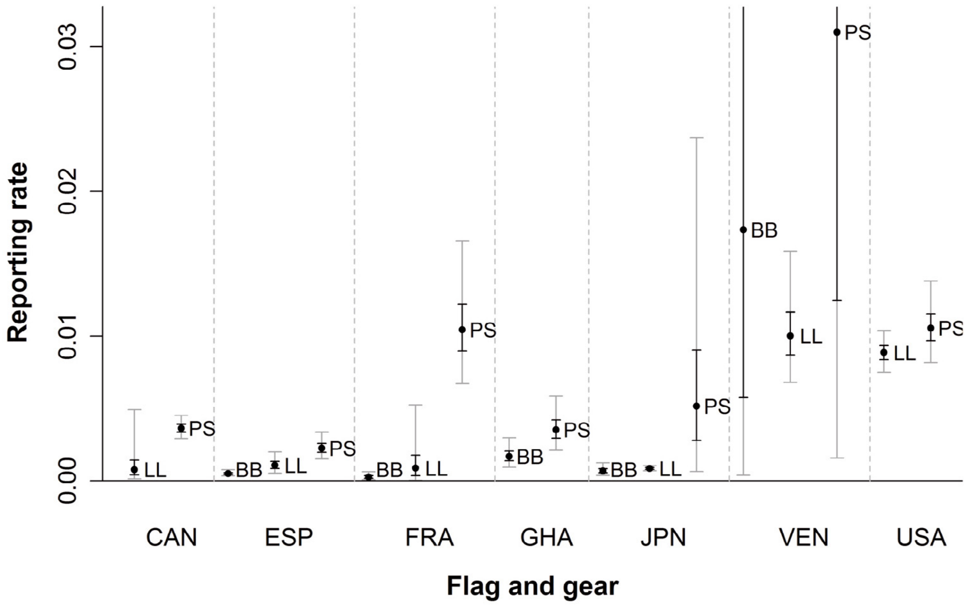

Fig. 2. A comparative plot of marginal posterior density estimates of reporting rate from the base case model based on biannual data that are not spatially disaggregated (returns before 91 days at liberty have been removed). The most extreme error bars of each point (in grey) represent the 95% probability interval. The smaller bars (in black) represent the 50% probability interval. The circular central points represent the median probability.

substantially lower than that estimated for fisheries elsewhere. Pollock et al. (2002a) for example, estimated reporting rates in the range of 0.24–0.95 in the Southern bluefin tuna fishery from 1991 to1995. Heifetz and Maloney (2001) estimated reporting rates in the range of 0.17–0.38 for the Alaskan sablefish fishery. The low estimated reporting rates for these Atlantic species raise serious questions over the cost efficacy of conventional tagging as a source of information for stock assessment.

While no figures have been explicitly stated for the opportunistic conventional tagging programs in the Atlantic, the Secretariat of the Pacific Community recently tagged 57 000 tuna at a cost of approximately US\$ $5{\times}10^6$ (SPC 2008). Accepting this as a guideline would put the total cost per conventional tag at around US\$ 100. At the low reporting rates estimated here (less than 1%), this equates to over US\$ 10 000 per recovery, putting conventional tagging in the same order of cost as other, more specialised, tagging methods. For example, where information regarding movement is required, the low reporting of conventional tags may provide sufficient justification for the use of more expensive technologies tailored to the collection of such information: such as pop-off satellite archival tags, which have much higher data recovery rates (greater than 85%, Sedberry and Loefer 2001; Block et al. 2005). Similarly, where information regarding exploitation is required, other technologies that do not rely on reporting by commercial fishers, such as genetic tagging (Palsboll 1999; Hoyle 2005) or the use of passive integrated transponder tags detected in the hold, codend or at the market (see Jorgensen et al. 2005) may prove more cost effective than conventional tagging (although this method remains unused in exploitation rate estimation for Atlantic tuna, billfish and shark species). This central issue of information value is a critical priority for management. ICCAT for example, are in the process of finalising a €19 × 10^6 research plan for bluefin tuna, planned to include an extensive conventional tagging program (Anon. 2009). Rather than the approximate calculations made here, an ongoing priority is a detailed evaluation of different tagging approaches in order to determine how best to improve precision and accuracy in the estimation of the quantities that are central to management, such as exploitation rate, abundance and movement rates.

In the absence of prior information regarding reporting rates, it is not clear how stock assessments may incorporate tagging data to provide exploitation rates and abundance. Reporting rates may have to be assumed constant among fleets, possibly on the basis of experiments undertaken in other regions and different species. The results of this paper suggest that such an assumption is likely to introduce varying magnitudes and directions of bias among the exploitation rate estimates of fleets that catch tuna, sharks and billfish in the Atlantic, generating conflicting data in an integrated assessment.

While the method of this study relies on extrapolation in time and space to estimate reporting rates for many fleets, three reporting rate estimates, those of the USA, Japanese and Venezuelan longline fleets, are estimated on the basis of fewer assumptions (they overlap to a greater extent with the observer fleet and are likely to have more similar gear selectivity). These reporting rate estimates were also less sensitive to the spatio-temporal resolution at which tag return rates are compared. Further work could use more disaggregated data to quantify reporting rates for these fleets at finer scales, and to investigate variability among vessels.

The estimates of purse seine reporting rate tend to increase as minimum days at liberty decrease. A probable explanation for this is that tags are often released from purse seine operations and these tags are caught soon after release in the same

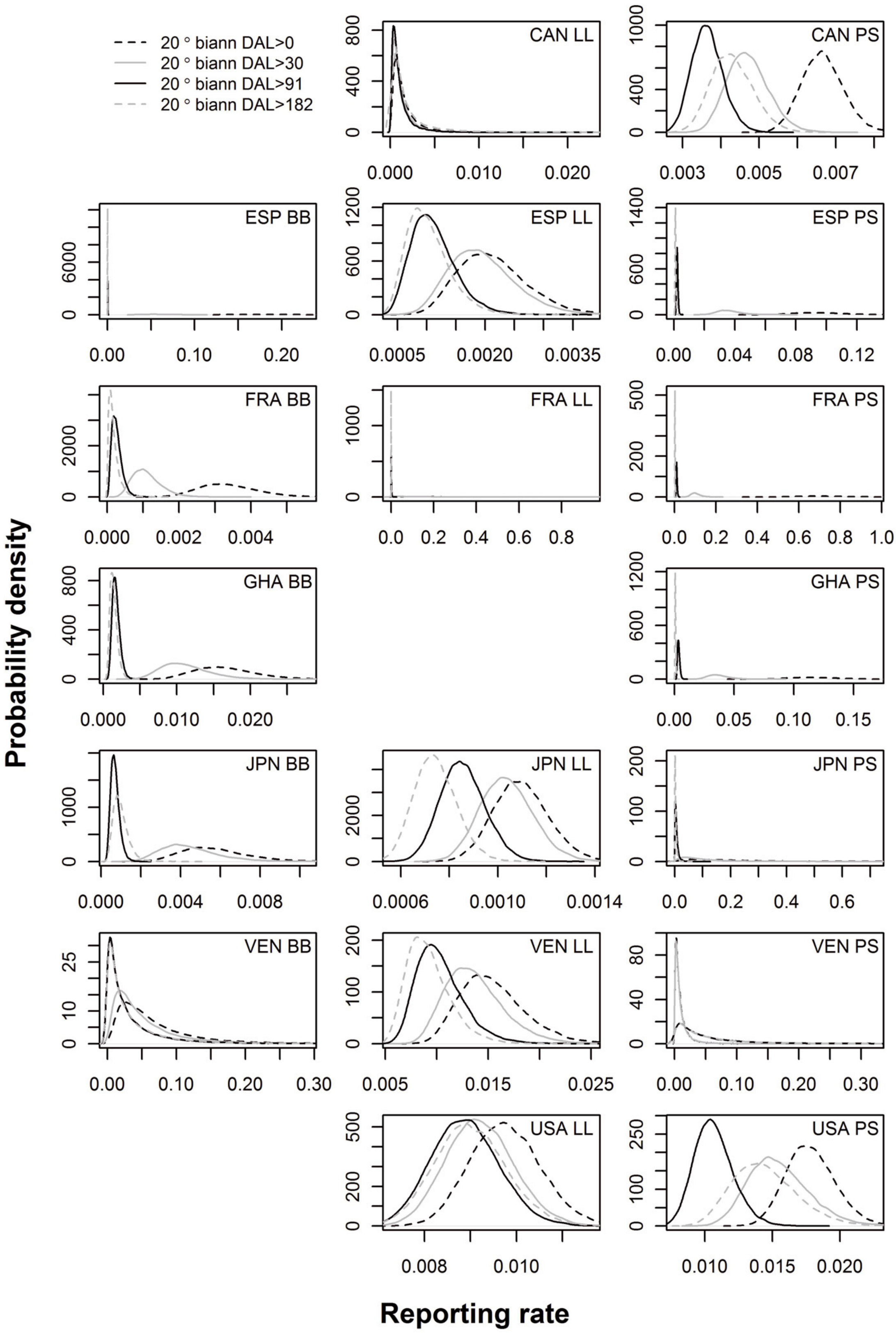

Fig. 3. Marginal posterior density plots of reporting rate illustrating the sensitivity of the return data to the minimum number of days at liberty (DAL). The base case data are represented by the solid black line. Durations of zero, 30, 91 and 182 are plotted here. All of these comparisons were made at the spatial scale of 20 × 20 degree ocean square and disaggregated biannually ("20° biann").

Table 5. Means and variance-covariance matrix for posteriors of logit-transformed reporting rate estimates.

	CAN LL	**CAN PS**	**ESP BB**	**ESP LL**	**ESP PS**	**FRA BB**	**FRA LL**	**FRA PS**	**GHA BB**
Mean	**–6.7281**	**–5.6094**	**–7.6016**	**–6.7958**	**–6.0735**	**–8.2058**	**–6.6296**	**–4.5218**	**–6.3453**
CAN LL	**0.8744**	0.0009	0.0000	0.0040	0.0021	0.0023	–0.0007	0.0037	0.0025
CAN PS		**0.0124**	0.0004	0.0001	0.0026	0.0012	0.0007	0.0024	0.0027
ESP BB			**0.0413**	0.0020	0.0056	0.0070	0.0417	0.0066	0.0061
ESP LL				**0.1190**	0.0026	–0.0002	0.0004	0.0023	0.0030
ESP PS					**0.0425**	0.0158	0.0033	0.0345	0.0325
FRA BB						**0.3106**	0.0083	0.0176	0.0183
FRA LL							**1.7531**	0.0043	0.0027
FRA PS								**0.0539**	0.0391
GHA BB									**0.0882**
GHA PS									
JPN BB									
JPN LL									
JPN PS									
USA LL									
USA PS									
VEN BB									
VEN LL									
VEN PS									

	GHA PS	**JPN BB**	**JPN LL**	**JPN PS**	**USA LL**	**USA PS**	**VEN BB**	**VEN LL**	**VEN PS**
Mean	**–5.6105**	**–7.2489**	**–7.0626**	**–4.9649**	**–4.7145**	**–4.5296**	**–3.3576**	**–4.5593**	**–2.8751**
CAN LL	0.0035	0.0007	0.0062	–0.0001	–0.0027	–0.0001	–0.0047	0.0024	0.0012
CAN PS	0.0024	0.0030	0.0016	0.0022	0.0010	0.0001	0.0015	0.0002	0.0039
ESP BB	0.0070	0.0044	0.0019	0.0059	0.0013	0.0002	0.0015	0.0001	0.0014
ESP LL	0.0017	0.0010	0.0009	0.0014	–0.0031	0.0011	0.0027	0.0011	0.0009
ESP PS	0.0331	0.0283	0.0035	0.0400	–0.0007	0.0010	–0.0018	0.0003	0.0101
FRA BB	0.0183	0.0139	0.0018	0.0146	0.0015	0.0004	–0.0002	0.0002	0.0043
FRA LL	0.0063	0.0030	0.0010	0.0062	–0.0092	–0.0006	–0.0060	–0.0013	0.0017
FRA PS	0.0418	0.0254	0.0039	0.0336	–0.0037	0.0011	–0.0006	0.0003	0.0090
GHA BB	0.0406	0.0291	0.0036	0.0329	–0.0008	0.0011	0.0013	0.0007	0.0100
GHA PS	**0.0690**	0.0272	0.0038	0.0325	–0.0024	0.0010	–0.0020	0.0004	0.0095
JPN BB		**0.0961**	0.0030	0.0268	–0.0019	0.0004	–0.0029	0.0004	0.0132
JPN LL			**0.0116**	0.0036	–0.0005	0.0011	0.0016	0.0002	0.0016
JPN PS				**0.8501**	0.0022	0.0025	0.0093	0.0001	0.0088
USA LL					**2.5642**	–0.0009	0.9411	0.0002	–0.0019
USA PS						**0.0477**	–0.0003	0.0003	0.0003
VEN BB							**1.8141**	0.0009	–0.0013
VEN LL								**0.0070**	0.0004
VEN PS									**0.0182**

fish aggregations. Estimates of reporting rate in purse seine fisheries also tend to be larger than those of the longline fleets of the same nation. In addition to the shoaling effect identified above, this phenomenon can be explained by differences in gear selectivity that we do not account for in this study. Purse seine gears tend to select for younger fish, which are expected to have higher mark rates due to loss of tags over time from tag induced mortality and tag shedding. These sensitivities may also be attributed to the scarcity of the reported tags. Increasing the minimum number of days at liberty creates more comparisons where Japanese and Spanish baitboat return rates are zero, which strongly reduces posterior estimates of their reporting rates.

It is not clear why the pelagic observer program has a higher return rate of tags that have been at liberty for longer than commercial fleets (in contrast, commercial reporting rate estimates decrease with increasing minimum duration at liberty). Due to natural mortality, tag induced mortality and tag shedding, it is expected that all fleets catch fewer tags that have been at liberty for longer. These tags may also be less likely to be caught in groups due to tag mixing over time. It may be the case that commercial operators detect or report fewer tags when return rates are low, while detection rate of the observer fleet remains relatively constant.

The method we employ in this study may be improved by additional observer data from other fleets and the use of the latest revision of the ICCAT conventional tagging database. The model can easily be extended to include archival tag data where available. In addition, tag seeding experiments can be used to describe priors for reporting rates, thereby providing the network of mark rates and other reporting rate estimates. Despite the reliance on several assumptions, this analysis is the first to estimate reporting rates of conventional tags for specific flags and gears fishing the Atlantic. By providing probability density functions and covariance matrices for the estimated reporting rates, uncertainty over these may be accounted for in subsequent analyses.

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