

Running head: Estimating historical minke whale bycatch in Japan

## Partitioning variance components and reporting rates to estimate historical bycatch: an example for minke whales in Japan

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**Abstract:**

Bayesian hierarchical modeling provides a convenient framework for extending generalized linear models to situations in which the model inputs consist of multiple information types, mixed sets of assumptions, different variance structures, and so on. We demonstrate a unique example, using historical bycatch data for minke whales around Japan, of how the flexibility of Bayesian hierarchical modeling permitted us to estimate parameters and parameter uncertainty that would have otherwise been confounded or cumbersome to estimate. The datasets consist of reported minke bycatch and fishing effort (nets) from 1979 – 2009. Over this time period, reported bycatch has increased. The questions are whether this has been due to increases in catch or increases in reporting rate, and what the historical bycatch levels may have been. Information about reporting history suggests that reporting rate was  $\ll 100\%$  prior to 2001 but may have been comparatively complete since 2001. This knowledge allowed us to use data from 2001 to 2009 to estimate process variance and bycatch trend parameters. Conditional on these estimates, sampling variance and reporting rate trend parameters are estimated from 1979 – 2000. All parameter estimation is done within a single framework, providing Bayesian posterior distributions for all parameters that account for all uncertainty and covariance in other parameter estimates. The results of our analysis suggest that reporting rate trend was more important than BPUE trend in explaining the trend in reported bycatch although both appear important. Model estimates fit the data well and substantially better than a model that did not estimate a reporting rate trend. Historical bycatch point estimates (Bayesian posterior means) from the model ranged from 28 in 1979 to 92 in 2000, with substantial uncertainty in the earliest years of the time series (e.g., CV up to 0.54) that transparently documents our poorer knowledge as we look further back in time. Summarizing the Bayesian posterior distribution for mean annual bycatch, the average number of minke whales taken incidentally from 1979 – 2009 was estimated to be between 55 and 101 per year (95% credible interval), with a posterior mean estimate of 76. This summary could be useful for setting bounds of annual bycatch for use in the Implementation Simulation Trials (ISTs), but we acknowledge that there may be other sources of uncertainty, particularly with the data prior to 2001, that cannot be taken into account.

**Keywords:** Asia, Fisheries, Gillnets, Incidental catches, Modelling, Statistics, Trends

**INTRODUCTION**

The IWC's Revised Management Procedure (RMP) requires a complete time series of catches, including incidental takes, that can be allocated to each cetacean stock that may potentially be exploited. For any population subject to exploitation, there may be several different ways, or variants, in which the RMP could be applied based on different hypotheses about stock structure and the spatial distribution of different stocks. Implementation Simulation Trials (ISTs) are used to investigate the performance of different variants across the range of scenarios that is considered plausible, including uncertainty in total historic and future incidental catches (Punt and Donovan, 2007). For the purposes of ISTs conducted as part of the RMP, it is not necessary to agree on best estimates but trials need to be conducted using values that cover what is considered to be the plausible range. For calculating actual catch limits, point estimates would be required.

The purpose of the current analysis is to describe a method that could provide improved point- and plausible range estimates for historical time series of whale bycatch in some data circumstances. We developed the approach for analysis of reported minke whale bycatch data around Japan. Incidental take of North Pacific minke whales around Japan occurs in small-scale trap net fisheries (in waters  $< 27\text{m}$  depth) and large-scale trap net fisheries set in deeper waters. Large-scale nets account for most minke whale bycatch; these nets include seasonally used salmon nets in Northern Japan (Aug – Nov) and other large-net operations that operate year-round in various areas (Tobayama et al. 1992).

**Historical Context**

In 1999, the IWC Scientific Committee agreed that, for the purpose of conducting ISTs, an appropriate range of annual incidental take estimates of North Pacific minke whales by Japan was 25-75 (IWC, 2000). At that time there were two existing estimates for minke bycatch in Japan: an average annual take of 93 during the 1980s based on independent observations of around 20 set nets extrapolated to the whole country (Tobayama *et al.*, 1992), and a revised estimate of 57 based on reanalysis of the same data by Inoue and Kawahara (1999). The latter estimate corrected some errors of the previous analysis but it was suggested by Kasuya that there may have been an overcorrection for some factors (IWC, 2000). He questioned the assumption by Inoue and Kawahara (1999) that small-type trap nets cannot catch minke whales and the allowance they made for some nets being seasonal. In

addition to these estimates, a time series of reported bycatch has been provided to the IWC by Japan since 1979 (Figure 1). For the period 1980-89 the average reported bycatch was 5 minke whales, increasing to 18 for the period 1990-1999. At the time it was stated that Japan ‘still believed that the official statistics were more reliable’ than either of the estimates (JCRM 2 p.84). But Tobayama *et al.* (1992) noted that only one of six minke whales caught by monitored nets during 1979-1987 and 1990 was officially reported, suggesting a significant negative bias in the reported takes.

Market surveys in Japan during 1993-1999 revealed a higher proportion of J-stock minke whales on the market (31%) than could be explained by reported bycatch and whaling. An incidental take estimate of 100 whales throughout Japan (J-stock + O-stock) was more consistent with the market proportions by stock (Baker *et al.*, 2000). Therefore, at the 2001 meeting it was suggested that the plausible range of bycatch around Japan be modified to span 25 – 100, to cover the entire range of values considered plausible by at least some members (IWC, 2002a). The option of including a single trial with baseline levels of 50 and 100 to test the reliability of interpolation/extrapolation with respect to bycatch was acceptable to all SC members except those belonging to the Japanese delegation (. Given the lack of consensus the chair ruled that the trials specification should remain the same as the previous range used, i.e., 25 – 75 (IWC, 2002a). The discrepancies between the official statistics and the independent estimates remained unresolved.

In 2001 new regulations (Ministerial Ordinance 92) were introduced in Japan allowing trap net fishermen to kill whales that are found in their traps, and to sell these on to the market, provided they register the animal by supplying a DNA sample (IWC, 2002b). This resulted in a large increase in the reported bycatch after July 2001 (Figure 1). This prompted new discussions in 2003 by the Committee as to the appropriate range of annual bycatch mortality to consider in the ISTs. The option of the lower level (annual average of 25) was retained but the option for the upper bound was raised from 75 to 100 on the basis that ‘the Chair determined that most members considered a value of 100 to be appropriate’ (IWC, 2003). The reported bycatch prior to 2001 is now believed to have been grossly under-reported (Lukoschek *et al.*, 2009; Hakamada and Ishikawa, 2010) and there is a need for further analyses in order to construct a valid time series estimate of historical bycatch.

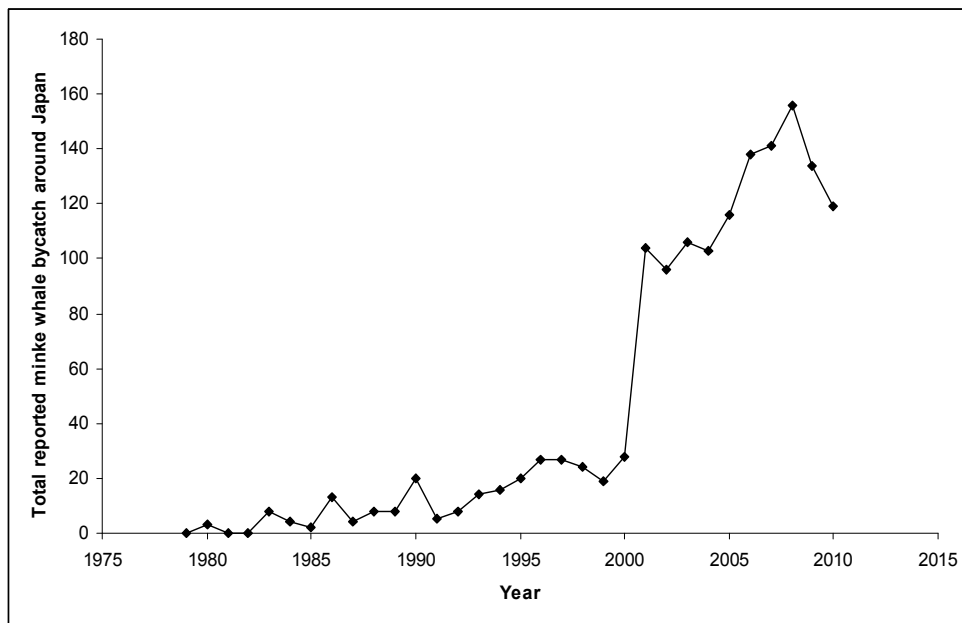


Figure 1. Reported bycatch of minke whales in Japan from annual progress reports to the IWC. Two regulatory changes (in 1990 and 2001) concerning sales and reporting of whale catches underlie to two pronounced changes in the time series: first, a reduced variance in reported values from 1991 – 2000 relative to earlier years; and second, a sharp increase in reported bycatch after 2000.

### Reconstructing historical time series of bycatch

The upper bound of 100 whales per year used in the 2003 IST does not take into account historical variation in the minke whale bycatch per unit effort (BPUE), fishing effort, or reporting rates that underlie variation

in reported (Figure 1) and ultimately true bycatch. Variation in BPUE could occur as a result of changes in minke whale abundance and/or fishing practices (e.g., fishing locations, net sizes, time spent fishing) such that the same level of reported fishing effort (reported simply as ‘number of nets’; Table 1) actually catches more or fewer whales.

Hakamada and Ishikawa (2010) used a Generalized Linear Model (GLM) to estimate BPUE trends in an attempt to correct the reported bycatch numbers prior to 2001. Key assumptions for their analysis were:

- (1) all bycatches since 2001 have been reported;
- (2) bycatch rates are proportional to population size;
- (3) a constant fraction of bycatch was reported prior to 2001.

The latter two assumptions imply that the increasing trend in reported and estimated bycatch from 1979 - 2001 reflects increasing minke whale abundance during that time. Direct evidence for an increasing J-stock abundance trend is inconclusive and contradictory (Kitakado *et al.*, 2010; Song, 2011), and alternative hypotheses may also explain increasing BPUE, including increasing trends in reporting rate, changes in actual fishing methods and locations, shifts in minke whale distribution patterns, or a combination of these factors.

Data are generally not available to evaluate all these nuanced hypotheses. However, based on available time series data for reported bycatch and fishing effort reported as the number of nets used per year (Table 1), and the assumption that bycatch reporting since 2001 has been more or less complete, we show that it is possible to tease apart the contribution of trends in reporting rate vs. trends in BPUE as explanatory factors of the observed trends in reported bycatch. In doing so, we are able to construct a more realistic estimated time series of bycatch that provides a much better fit to the data than does the model of Hakamada and Ishikawa (2010), and we make no assertion about the relationship between BPUE trends and minke whale population trends. Our approach provides a clear description of the uncertainty in bycatch time series estimates, which based on our assumptions, can provide empirical upper and lower bounds of bycatch that could be used in the ISTs. The assumption of perfect bycatch reporting since 2001 is difficult to validate; however, the average reported bycatch of 123 whales since 2001 is higher than the independent estimates from the 1990s based on observation and market sampling (Tobayama *et al.*, 1992; Inoue and Kawahara, 1999; Baker *et al.* 2000). Thus there is no direct evidence for substantial under-reporting since the new regulations in 2001 came into place, and so the assumption that reported numbers since 2001 are not substantially biased may be reasonable enough for purposes of the analysis.

Table 1. Data used to fit model and estimate bycatch in Japan from 1979 – 2000 (net data from Ministry of Agriculture, Forestry and Fisheries, Japan in Tobayama *et al.* (1992) and Hakamada and Ishikawa (2010)). Bold italicized net values are our estimates, not reported values. Estimates between 1971 and 1993 were interpolated from a simple quadratic model fit to reported effort between 1970 and 1994 ( $n = 6$ ); uncertainty in fitted values was propagated in the Bayesian bycatch estimation model. Estimates after 2006 were fixed at the 2006 reported value.

Year	Reported bycatch	Large-scale trap nets (incl. salmon nets)
1970		1210
1971		<b>1210</b>
1972		<b>1248</b>
1973		<b>1284</b>
1974		<b>1319</b>
1975		1272
1976		<b>1387</b>
1977		<b>1419</b>
1978		<b>1450</b>
1979	0	<b>1479</b>
1980	3	1544
1981	0	<b>1535</b>
1982	0	<b>1562</b>
1983	8	<b>1591</b>
1984	4	<b>1614</b>
1985	2	1646
1986	13	<b>1662</b>
1987	4	<b>1681</b>

1988	8	<b>1702</b>
1989	8	1742
1990	20	<b>1741</b>
1991	5	<b>1755</b>
1992	8	<b>1771</b>
1993	14	<b>1787</b>
1994	16	1782
1995	20	1755
1996	27	1740
1997	27	1704
1998	24	1701
1999	19	1702
2000	28	1692
2001	104	1689
2002	96	1672
2003	106	1655
2004	103	1633
2005	116	1607
2006	138	1589
2007	156	<b>1589</b>
2008	134	<b>1589</b>
2009	119	<b>1589</b>

## METHODS

The GLM used by Hakamada and Ishikawa (2010) has the form:

$$\log[E(c_t)] = \log(n_t) + \beta_0 + \beta_1 t + \beta_2 r_t, \quad (1)$$

where  $E(c_t)$  is the expected level of bycatch in year  $t$ ,  $\log(n_t)$  is an offset term for the number of fishing nets at  $t$ , and  $r_t$  is a dummy variable indicating whether year  $t$  is pre-2001 ( $r_t = 1$ ) or is during the period 2001-2006 ( $r_t = 0$ ). These two time periods correspond to different bycatch reporting-rate periods; reporting rate is assumed to be 1 from years 2001 onward, whereas the pre-2001 rate must be estimated. The parameters  $\beta_0$  (intercept),  $\beta_1$  (trend in BPUE), and  $\beta_2$  (pre-2001 reporting rate) are to be estimated by the GLM.

We begin with a simple extension of equation 1, to more explicitly address sources of variation in reported bycatch. First, we note that expected reported bycatch in year  $t$  reflects a deterministic component plus a stochastic component (i.e., some real year to year variation in BPUE),  $\eta_t$ , and also sampling error due to variation in annual reporting rate,  $\tau_t$ . Thus,

$$\log[E(c_t)] = \log(n_t) + \beta_0 + \beta_1 t + \beta_2 r_t + \eta_t + \tau_t. \quad (2)$$

Both error terms are assumed to be normally distributed with mean zero and variances  $\sigma_\eta^2$  and  $\sigma_\tau^2$ , respectively. Total variance equals  $\sigma_\eta^2 + \sigma_\tau^2$ . If we assume that process variance  $\sigma_\eta^2$  is relatively constant through time, and that reporting rate from 2001-2006 is 1 such that sampling error terms  $\sigma_\tau^2$  and  $\tau_t = 0$  during these years, then these variance components can be independently estimated. This is because variance in years 2001 – 2006 provides a direct estimate of  $\sigma_\eta^2$ , and then  $\sigma_\tau^2$  in 1994 – 2000 equals total variance minus  $\sigma_\eta^2$ . Treating the recorded bycatch data,  $c_t$ , as a Poisson random variable (with possible overdispersion accounted for by the additional variance terms), the data likelihood is given by:

$$c_t \sim \text{Pois}\{E(c_t)\},$$

$$\text{where } E(c_t) = \exp\{\log(n_t) + \beta_0 + \beta_1 t + \beta_2 r_t + \eta_t + \tau_t\}.$$

Then, expected *actual* (vs. reported) bycatch,  $b_t$ , may be estimated each year as:

$$E(b_t) = \exp\{\log(n_t) + \beta_0 + \beta_1 t + \eta_t\}. \quad (3)$$

As previously noted, steady observed increases in reported bycatch from 1979 to 2000 (i.e., prior to complete reporting of bycatch) could reflect increasing BPUE (as assumed in model forms above), increasing reporting rates (e.g., with more awareness about the issue), or both. To estimate trends in both BPUE and reporting rates, we extend the model further:

$$\log[E(c_t)] = \log(n_t) + \beta_0 + \beta_1 t + \beta_2 r_t + \beta_3 \cdot t \cdot r_t + \eta_t + \tau_t. \quad (4)$$

Compared to equation 2, this model has new parameter,  $\beta_3$ , which represents trend in reporting rate during 1979 – 2000. Analogous to our arguments for being able to estimate both variance components (process and sampling error) in the model, our ability to estimate both trend parameters ( $\beta_1$  and  $\beta_3$ ) is conditional on the assumption that reported bycatch from 2001 – 2009 is relatively error free. This enables us to assume no reporting rate trend during this time period ( $\beta_3 \cdot t \cdot r_t = 0$ ), so that the 2001 – 2009 data provide a direct estimate of  $\beta_1$  (equation 3). Estimates of  $\beta_3$  for the 1979 – 2000 dataset are then largely conditioned on the estimate of  $\beta_1$  from the 2001 – 2009 dataset, given the additional assumption that the mean trend in BPUE has remained constant through time (e.g., average annual increase in BPUE throughout the 1980s was similar to average annual increase in BPUE during the last decade).

As a final modification of our model, we allow reporting rate variance  $\sigma_\tau^2$  to be estimated separately for the periods 1979 – 1990 ( $\sigma_{\tau 1}^2$ ) and 1991 – 2000 ( $\sigma_{\tau 2}^2$ ). The basis for this comes from regulation changes in Japan concerning whale reporting in 1990. In essence:

$$\log[E(c_t)] = \begin{cases} \log(n_t) + \beta_0 + \beta_1 t + \beta_2 + \beta_3 t + \eta_t + \tau_{1,t} & \text{for } t < 1991; \\ \log(n_t) + \beta_0 + \beta_1 t + \beta_2 + \beta_3 t + \eta_t + \tau_{2,t} & \text{for } 1991 \leq t \leq 2000; \\ \log(n_t) + \beta_0 + \beta_1 t + \eta_t & \text{for } t > 2000. \end{cases} \quad (5)$$

Parameters were estimated using a Bayesian MCMC approach in WinBUGS 1.4.3 (code available from authors). We do not provide a general background on the advantages of Bayesian analyses here, as numerous overviews have already been published in the ecological literature (e.g., Wade, 2000; Ellison, 2004; Cressie *et al.*, 2009). We note, however, that Bayesian methods are well suited to the current problem because of their convenient handling of variance partitioning, multiple data types, error propagation, and the explicit inference they provide about processes of interest. In our case, this means that uncertainty in estimates of interpolated fishing effort (Table 1) and of parameters based on data from years  $\geq 2001$  were fully incorporated into conditional estimation of other parameters, and that uncertainty in parameters shared across all time periods allowed for convenient estimation of parameters unique to specific time periods (e.g.,  $\sigma_{\tau 1}^2$  and  $\sigma_{\tau 2}^2$ ). More importantly, it means we can isolate certain parameters to explicitly estimate the posterior distribution for actual bycatch (equation 3) in a way that takes all parameter uncertainty and estimation covariances into account. We specified vague prior distributions for all estimated parameters (uniform [0,100] for  $\sigma$  terms, and normal [0, large] for  $\beta$  terms). We summarize posterior distribution summaries of parameters constructed of 200,000 MCMC samples (2 chains x 100,000 samples), thinned by 2 to 100,000 samples, following a burn-in period of 100,000 samples on each chain.

## RESULTS

Parameter estimates (posterior distribution summaries) for the full model (equation set 5) are in Table 2. These results indicate strong evidence of increasing trend in both in BPUE (avg rate = 0.056 increase per year) over the full period 1979-2009 and reporting rates over the period 1979-2000 (avg rate = 0.09 increase per year), with reporting probability increasing at a faster rate than BPUE. We reiterate that the BPUE trend estimate is based largely on post-2000 data, with the pre-2001 reporting rate trend estimate conditioned largely on the BPUE estimate (given the assumption that a single BPUE trend can be applied to the entire time series). A majority of the variance in the pre-2001 reportings can be attributed to sampling error rather than process variance. Sampling error, i.e., reporting rate variance, appeared particularly high prior to 1990.

Table 2. Posterior distribution summaries for model parameters, fit to 1979 – 2009 data. Note that  $t = 1$  implies year 1970, the first year in which effort data were reported (see Table 1). CRI = Bayesian credible interval.

Parameter	Mean	SD	95% CRI
$\beta_0$	-4.638	0.745	-6.503, -3.332
$\beta_1$ (BPUE)	0.056	0.021	0.020, 0.108
$\beta_2$ (Pre-2001 effect on reporting rate)	-3.823	0.963	-5.681, -1.799
$\beta_3$ (Pre-2001 trend in reporting rate)	0.090	0.031	0.027, 0.152
$\sigma_\eta$ (Process variance)	0.108	0.059	0.013, 0.248
$\sigma_\tau$ (Reporting rate variance, 1991-2000)	0.189	0.137	0.009, 0.516
$\sigma_\tau$ (Reporting rate variance, 1979-1990)	0.873	0.389	0.335, 1.809

Our estimates of reporting rate trends result in historical bycatch estimates that are generally higher than those of Hakamada and Ishikawa (2010) (Table 3). The earliest estimates are highly uncertain, however, appropriately reflecting lower reporting rates (including reported values of zero) and higher reporting rate variance in earlier years. The posterior mean for average annual bycatch from 1979 – 2000 was 58 animals, with 95% upper and lower credible interval bounds of 29 and 93. For the full period 1979-2009, the posterior mean is 76 animals per year, with 95% upper and lower bounds of 55 and 101. Some version of these estimates (year-specific, or an annual rate summary across years) could potentially inform the IST process.

Our model estimates appear to fit all the data well, for both the *reported* bycatch dataset (1979 – 2000), which is known to be a biased representation of actual bycatch (Figure 2a), and for the dataset from 2001 onward, which is assumed to accurately represent actual bycatch (Figure 2b). In contrast, when we fit a model in which reporting rate trends were not estimated (see equation 2), it fits portions of the dataset poorly (Figure 3). The estimates by Hakamada and Ishikawa (2010) come from this latter type of model.

Table 3. Reported bycatch, estimates as reported by Hakamada and Ishikawa (H & I, 2010), and estimates obtained by our model, from 1979 – 2000. We provide Bayesian posterior means (point estimates), standard deviation and CV of the estimate, and the 95% Bayesian credible interval estimates.

Year	Reported	H & I	Our Model Estimates			
			Mean	SD	CV	95% CRI
1979	0	0	28	15	0.54	6, 65
1980	3	11	32	16	0.50	8, 72
1981	0	0	32	16	0.50	8, 71
1982	0	0	34	16	0.47	9, 73
1983	8	30	38	18	0.47	11, 81
1984	4	15	40	18	0.45	12, 81
1985	2	8	42	17	0.40	13, 82
1986	13	49	46	19	0.41	16, 91
1987	4	15	48	18	0.38	17, 89
1988	8	30	51	19	0.37	20, 94
1989	8	30	55	19	0.35	22, 98
1990	20	75	59	19	0.32	26, 104
1991	5	19	59	18	0.31	25, 97

1992	8	30	63	18	0.29	29, 102
1993	14	53	70	19	0.27	35, 111
1994	16	60	73	18	0.25	39, 114
1995	20	75	77	18	0.23	44, 117
1996	27	102	83	18	0.22	51, 125
1997	27	102	84	17	0.20	53, 123
1998	24	91	85	16	0.19	54, 118
1999	19	72	85	15	0.18	53, 114
2000	28	106	92	14	0.15	62, 121
Mean <sub>1979-2000</sub>	12	44	58	16	0.28	29, 93
Mean <sub>1979-2009</sub> <sup>a</sup>	43	66	76	11	0.15	55, 101

<sup>a</sup> See Table 1 for reported values from 2001 – 2009, which are assumed to be accurate bycatch values

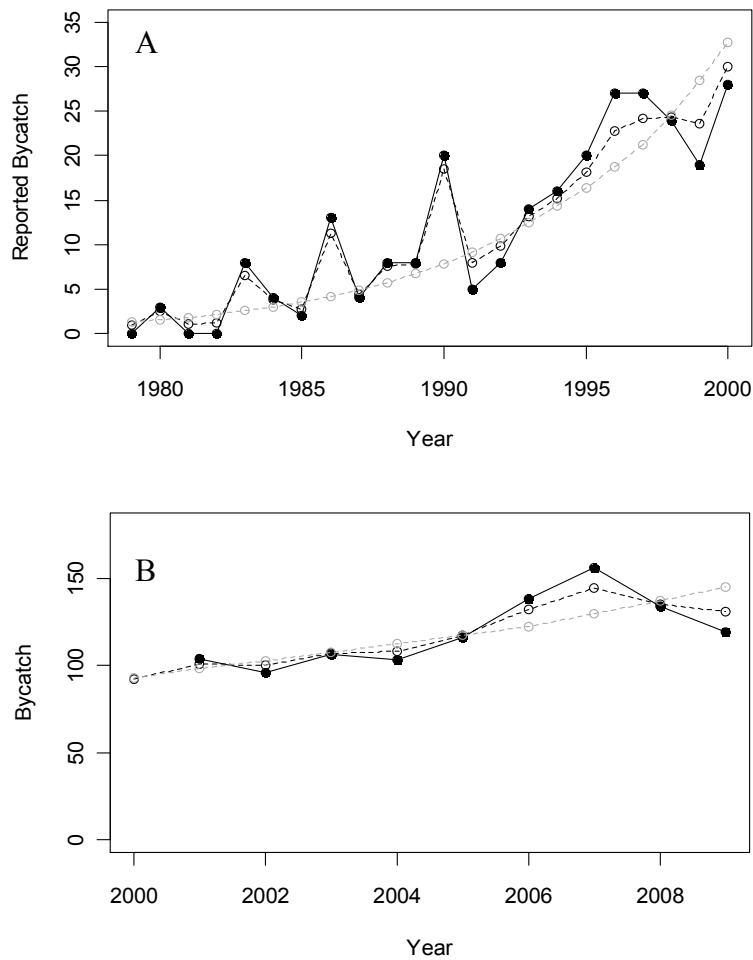


Fig 2. Results of full model incorporating BPUE and reporting rate trends and time-dependent sampling variance (equation set 5). Both panels show reported bycatch data (solid points), model fitted estimates that reflect process variance estimates (open black points and dotted line), and the mean fitted trend (open gray circles with dotted line). Top panel (A) shows years 1979-2000, when bycatch reporting rate was  $< 1$ , such that the model estimates depict fitted values for *reported* bycatch, not actual bycatch estimates. Bottom panel (B) shows actual bycatch estimates and data for 2001 onward, when reporting rate was assumed to be 1.



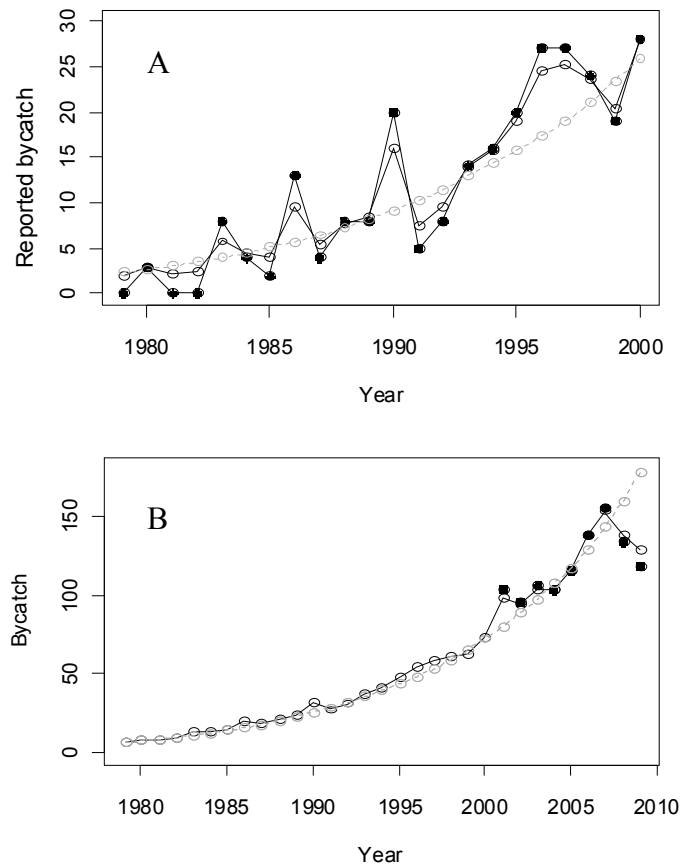


Fig 3. Results of model incorporating BPUE but not reporting rate trends (equation 2). Both panels show reported bycatch data (solid points), model fitted estimates that reflect process variance (open black points), and the mean fitted trend (open gray circles with dotted line). Top panel (A) shows years 1979-2000, when bycatch reporting rate was  $< 1$ , such that the model estimates depict fitted values for *reported* bycatch, not actual bycatch estimates. Bottom panel (B) shows actual bycatch estimates in all years (open points), but only shows data (solid points) for 2001 onward, when reporting rate was assumed to be 1. Note the consistent underfit of the mean trend to the data during the mid 1990s (A) and the overfit of the mean trend to the data at the end of the time series (B).

## DISCUSSION

Our analysis exemplifies the flexibility provided by Bayesian specification of generalized linear modeling for dealing with ‘messy’ time series datasets. The time series of reported minke whale bycatch in Japan reflected some known changes in variance sources and reporting rates across different time periods. By defining these explicitly in the model, we were able to account for these factors to estimate trends in both reporting rates and BPUE that satisfactorily describe the reported bycatch dataset and suggest an increase since 1979 in true bycatch of minke whales around Japan. Our BPUE trend point estimate (0.056 increase per year) is quite similar to the abundance trend point estimate for J-stock whales in Korea over the time period 2000 – 2009 (0.049 increase per year) (An *et al.*, 2010a; Song, 2011); this might provide some support for the hypothesis that recent BPUE trends are related to minke whale population size. A rough comparison of J-stock abundance estimates between 1983 (7100 – 10400; IWC, 1984) and 2009 (16,200; Kitakado *et al.* 2010) also suggests possible J-stock abundance increases, although the earlier effort-based estimates (e.g., CPUE, SPUE) are rather uncertain and may have only estimated stock abundance in a portion of its range. On the other hand, the abundance trend estimate by Song (2011) was not statistically significant, and Kitakado *et al.* (2010) found no evidence of abundance trends for the J-stock over the same time period when analyzing more recent abundance estimates. Further, J-stock bycatch data from Korea (1996

– 2008, An *et al.*, 2010b) does not contain any trend signal. In short, evidence of increasing minke whale abundance remains equivocal, and the explanations for apparent increases in BPUE remain uncertain.

Although our case study describes a specific analytical solution for dealing with the particular circumstances of the minke whale dataset, our approach to the problem may have some general applicability to other study systems because the type of issues we addressed here are probably not unique to this dataset. That is, many historical time series of reported catches or bycatches are likely to reflect periodic changes in management or policies that result in non-constant sources of variation and bias in the dataset, and for which variation in reporting rate vs. actual bycatch may be difficult to tease apart. Our analysis provides an example of how models may be constructed to explicitly account for sources of variation that are tied to known factors such as management changes and how to take advantage of portions of the dataset to estimate parameters that might otherwise seem confounded.

Our estimates of historical minke whale bycatch in Japan are generally consistent with previous estimates, considering the range of uncertainty in all estimates. Baker *et al.* (2000) estimated that approximately 100 minke whales needed to be incidentally killed per year (from 1995 – 1998) in Japanese coastal net fisheries to explain the estimated proportion (0.31) of J-stock samples found in Japanese markets the mid-late 1990s. This is higher than our bycatch point estimates for this time period (77 - 85), but well within our corresponding credible interval estimates (Table 3). Moreover, the uncertainty in the proportional contribution of J-stock whales to market products (0.19 – 0.43, Baker *et al.* 2000) translates into wide confidence intervals of the number of animals incidentally taken per year, and their estimates. A revised estimate for the market proportion of J-stock whales from 1997 – 2004 (0.46, 95% CI: 0.38 – 0.56), which was conditioned on the null hypothesis of no historical trend the proportions, generated annual Japanese bycatch estimates spanning 67 – 260 per year from 1997 – 2001 applied across a range of analysis scenarios (Lukoschek *et al.* 2009). The lower portion of this range (*< ca.* 120) is consistent with our credible interval estimates for these years (Table 3). Tobayama *et al.* (1992) extrapolated an estimate of 93 animals per year during the 1980s; this is considerably higher than our upper credible estimates for this time period, but Tobayama *et al.* (1992) noted that their estimation was likely biased because it was based on a small non-representative sample of trap nets. The revised estimate of 57/year by Inoue and Kawahara (1999), based on the same dataset used by Tobayama *et al.* (1992), is well within our distributions of estimates during the 1980s (Table 3). In short, all available estimates including ours contain substantial uncertainty and depend on important assumptions. Considering this, they tend to corroborate each other reasonably well.

Important caveats to our analysis are the assumptions that bycatch reporting rate since 2001 has been 100% and that the overall trend in reported takes can be described by a single mean trend in BPUE from 1979-2009 and a single mean trend in reporting rate from 1979-2000. The first assumption has already been acknowledged (see ‘Reconstructing historical time series of bycatch’). Although difficult to fully substantiate, the assumption of complete reporting since 2001 may be accurate enough that potential bias (i.e., underestimation of bycatch) may only be minor. The second assumption is more important. It seems improbable that either reporting rates or BPUE would increase more or less monotonically at the same rate over the course of 30 years, given the dynamics of whale populations and movements, management schemes, and fishing methods. That said, we note that trend analyses of time series data are typically and often necessarily conditional on such assumptions. Ideally, the model outputs are deemed useful in spite of their limitations, and by partitioning process errors and time-dependent sampling errors, our annual bycatch estimates may at least account for minor assumption violations. For example, in the mid-1990s, our mean trend estimates consistently underfit the data slightly (Fig. 2a), possibly because of some change in reporting rate or BPUE that was not captured by our model, but our explicit incorporation of annual sampling and process errors into the fitted estimates brought these more in line with the data, and the width of the credible intervals reflect the disparity between the mean trend and the observations.

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