Bering-Chukchi-Beaufort Seas bowhead whale (*Balaena mysticetus*) abundance estimate from the 2019 ice-based survey

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ABSTRACT

An ice-based visual survey of the Bering-Chukchi-Beaufort Seas stock of bowhead whales (*Balaena mysticetus*) was conducted in spring 2019 near Utqiaġvik (formerly Barrow), Alaska. A Horvitz-Thompson-type estimator is used to estimate population abundance from the resulting data, correcting for detection probabilities, whale availability within visual range, and whale passage during periods of missed effort. Analytical methods mirror those used for the 2011 survey as much as possible; however, unlike 2011, no simultaneous acoustic monitoring was conducted in 2019, so the availability correction factor had to be estimated from past years. The estimated abundance was 12,505 with 95% confidence interval of (7,994, 19,560) and a CV of 0.228. This estimated abundance is markedly lower than the 2011 estimate of 16,820, but the 2019 confidence interval wholly encompasses the 2011 interval. We do not interpret this finding as evidence of a decline for many reasons including: highly unusual ice conditions, an unusual migration route that was sometimes too distant from observers to detect whales, failure to conduct watch because of closed leads during the early weeks of the migration when numerous whales likely passed, an unusually short perch, and hunters' heavy use of powered skiffs near the observation perch which likely disturbed the whales during the survey. Furthermore, bowhead health assessment information for 2019 suggests that harvested bowheads did not exhibit obvious reductions in health condition, and aerial surveys in summer 2019 indicated high calf production. Despite the challenges of this survey, the estimate is adequate for use with the International Whaling Commission's management procedure and complies with the survey requirements of the Aboriginal Whaling Scheme.

KEYWORDS: ARCTIC, BOWHEAD WHALE, ABUNDANCE ESTIMATE, SURVEY – SHORE-BASED, WHALING – ABORIGINAL

INTRODUCTION

Each spring the Bering-Chukchi-Beaufort (BCB) Seas stock of bowhead whales (*Balaena mysticetus*) migrates northeast past Utqiaġvik (formerly Barrow), Alaska, into the Beaufort Sea. Those whales are hunted, in limited numbers, by Alaskan Native hunters residing in communities along the migration route (Stoker and Krupnik, 1993). Subsistence hunting quotas are established by the International Whaling Commission (IWC) in accordance with its Bowhead Strike Limit Algorithm (IWC, 2003a). A requirement of this process is the provision of abundance estimates for this stock at least once every 10 years. The last published estimate was 16,820 (CV = 0.052) for 2011 from an ice-based visual and acoustic survey near Utqiaġvik (Givens *et al.*, 2016).

In 2019, a new ice-based visual survey was conducted. As much as possible, the same methods were used as in past surveys (Clark *et al.*, 1986; George *et al.*, 2013; George *et al.*, 1995; Krogman *et al.*, 1989; Zeh *et al.*, 1993).

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Unlike 2011, no acoustical component was included. This means that there is no contemporaneous estimate of the proportion of the population that migrated beyond visual range of the survey perch. This problem can be overcome by estimating an availability correction factor from past surveys, as has been done in some previous years (Zeh and Punt, 2005) when acoustic data were similarly unavailable. The 2019 statistical abundance estimation closely replicates the 2011 analysis in all other respects.

FIELD METHODS AND DATA COLLECTION

The ice-based visual survey methods used for BCB bowheads have remained remarkably consistent since the late 1970s. Details are provided in previous reports (Clark *et al.*, 1986; George *et al.*, 2013; George *et al.*, 1995; George *et al.*, 2004; Krogman *et al.*, 1989; Zeh *et al.*, 1993). Briefly, a perch was constructed on an ice ridge or pressure ridge near the edge of the shore-fast ice adjacent to the shear zone where observers could scan the lead (i.e. the channel of open water); see Fig. 1. The lead may be entirely closed, ice-choked, narrowly open, or open many kilometres wide, depending on the wind and currents that shift the floating sea ice. A team of (usually) three observers atop the perch scanned the ocean for bowheads using the naked eye and binoculars. When bowheads were observed, the bearing, vertical angle and distance to the group were determined from theodolite readings, and the group size was recorded. Bowheads were tracked on nautical-type plotting paper to determine which animals were *not* previously sighted (i.e. 'New whales') and which animals were previously seen (i.e. 'Duplicates'). If an observer could not determine whether an animal was a New or Duplicate sighting, it was coded as 'Conditional', defined as having 50% chance of being previously sighted. A wide variety of other data were also collected about whale sightings (e.g. behaviour, swim direction, calves) and the environment (e.g. lead condition, visibility, time). A few whales were not located with the theodolite, and these required special treatment in the statistical analysis.

Although brief observations of conditions were made sporadically beginning on 16 March, the first full watch period was on 14 April once the ice and visibility conditions permitted. Numerous bowheads were seen on the 14 April watch. The perch was shut down on 23 May due to deteriorating and unsafe ice conditions. Between 14 April and 23 May, the perch was staffed as much as possible, with inactive periods depending on light, weather, ice conditions, and staff availability. For most of the season, we sought to maintain 18–24 hours of watch per day. Altogether, 630.65 hours of watch (whether qualifying or not; see below) were completed during this span

Fig. 1. Observers on the perch use the naked eye, binoculars, and theodolite to detect bowheads swimming past.

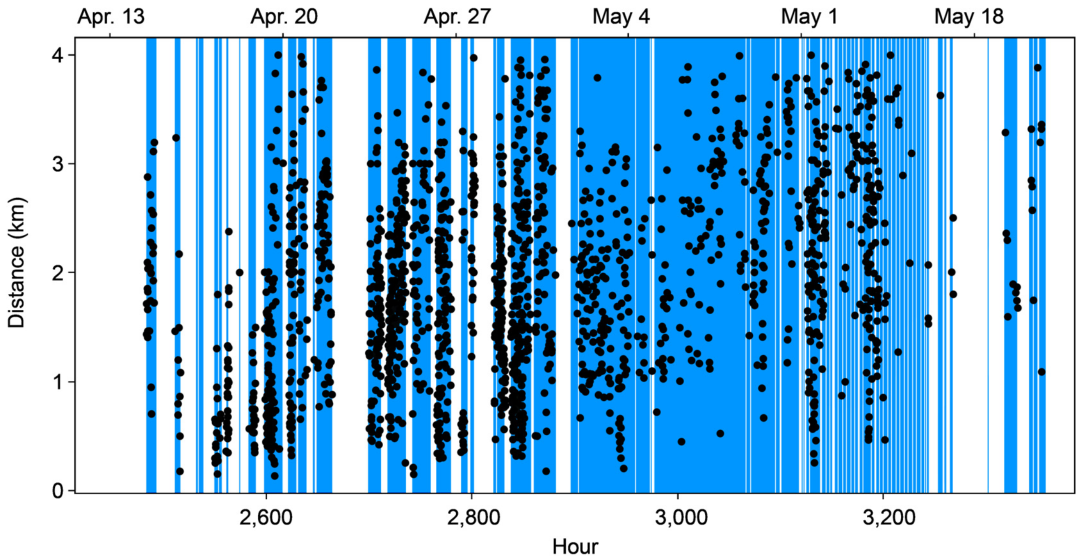

Fig. 2. Bowhead sightings made within 4 km of the perch during periods of qualifying watch effort (blue bands).

of 874.10 hours. We observed 2409 New and 353 Conditional whales; in 2011 these numbers were 3379 New and 632 Conditional in 787.50 total hours of watch. The 2019 counts include 34 New and 1 Conditional calves. For 2011, the calf counts were 13 New and 0 Conditional.

The statistical analysis is based on observations during qualifying conditions, specifically excluding times when the perch was not staffed, environmental conditions were deemed unsafe, the lead was closed and sea ice completely covered the survey area, or visibility was rated as poor or unacceptable. There were 544.77 hours of qualifying watch, during which 2018 New and 303 Conditional whales were seen within 4 km of the perch. For comparison, the 2011 abundance estimate was based on 530.02 hours of qualifying watch. Fig. 2 summarises the 2019 survey data. The blue shaded bands indicate periods of qualifying watch. Each point represents a New or Conditional sighting within 4 km of the perch. Sightings occurring during non-qualifying watch or beyond 4 km are not shown.

We use historical data to estimate a correction factor for availability. During some past surveys, including 2011, data were collected to directly estimate the proportion of whales passing within visual range (4 km) of the perch, denoted P_4. No such data were collected in 2019. Fig. 3 shows the estimated values of P_4 for past survey years plotted against the locations of the perch. The areas of the circles are proportional to the standard errors of the corresponding P_4 estimates, so years with smaller circles are known more precisely. The estimated P_4 values are taken from Zeh and Punt (2005), who derived them from either acoustic or aerial survey data or both, depending on the year. In 1985, three perches were used at different times, so we use a weighted average of the locations and P_4 estimates. The red triangle indicates the location of the perch in 2019.

STATISTICAL METHODS

Overview

As was done for past surveys, we analyse only New and Conditional sightings; Duplicates are not used. Previous abundance estimates have always treated a Conditional whale as half a whale, and we continue that approach here.

Let c_i denote the size of the *i*th sighted group, for $i = 1,\ldots,g$, where *g* is the number of groups sighted. Although grouping may vary during passage, we conceive the groups as being defined at the moment they pass the perch. Most (85%) groups were size 1, and the largest group was 6 whales.

We continue the past convention of treating 4 km as the limit of how far whales can reliably be seen from the perch. Our analysis assumes that bowheads are only available for sighting when they swim and surface within the 4 km radius visual detection zone. The sightings with estimated distances beyond 4 km are truncated from

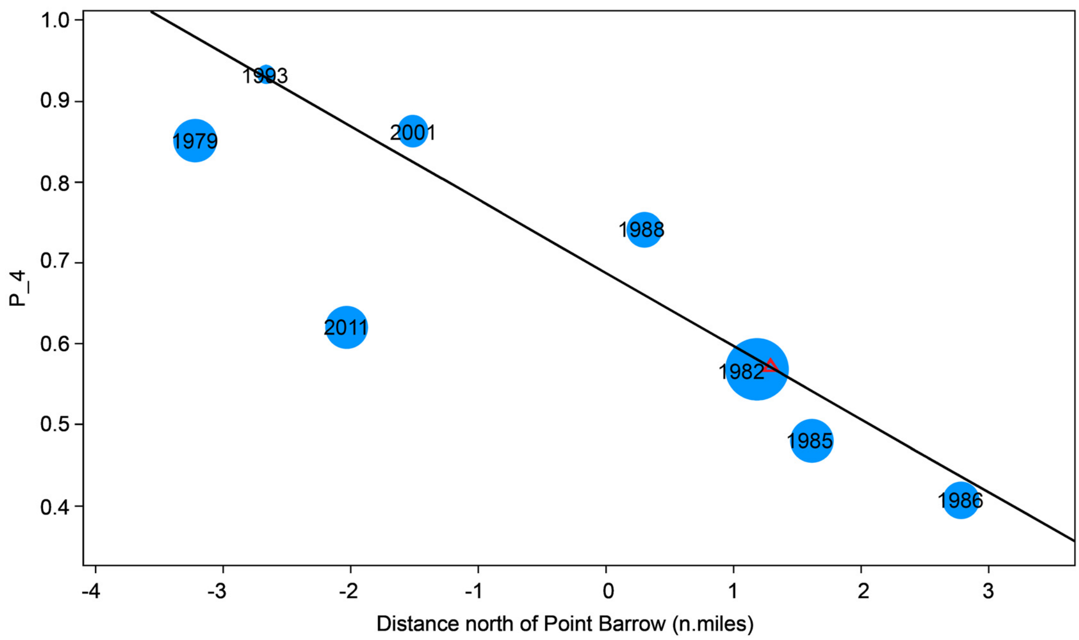

Fig. 3. Estimates of P_4 and the locations of the perch for past ice-based surveys. The areas of the circles are proportional to the standard errors of P_4 estimates. The red triangle indicates the location of the 2019 perch, for which no direct estimate of P_4 was obtained. Distance north of Point Barrow was calculated as the distance (north or south) in nautical miles from an E–W line at the latitude of the Point Barrow racon tower (71.3869 N).

the analysis. In some cases this occurred because of refraction; our theodolite locations suggested the whales were above the horizon. Let P_4 denote the proportion of groups that passed within 4 km of the perch. This is a correction factor for availability. Also, let p_i denote the detection probability, which is the conditional probability that the *i*th group was seen given that it was available.

Let N denote the unknown total population size. We use a scaled modified Horvitz-Thompson estimator (Borchers *et al.*, 2002; Horvitz and Thompson, 1952):

$$\hat{N} = \frac{1}{\hat{E}\hat{P}_4} \sum_{i=1}^{g} \frac{c_i}{\hat{p}_i} \tag{1.1}$$

where $1/\hat{E}$ is a correction for whales passing at missed times, and $\hat{P}_4$ and $\hat{p}_i$ estimate P_4 and p_i respectively. The correction factor $1/\hat{E}$ is estimated because-despite knowing the times when the perch was and was not operational – the number of whales that passed during those times is unknown because the passage rate varies during the season.

In their analysis of the 2011 survey data, Givens *et al*. (2016) used a similar estimator, namely

$$\hat{N}_{2011} = \frac{1}{\hat{E}} \sum_{i=1}^{g} \frac{c_i}{\hat{a}_i \hat{p}_i} \tag{1.2}$$

where $\hat{a}_i$ were sighting-specific corrections for availability. Our 2019 data does not permit estimation of $\hat{a}_i$ so we use a predicted value $\hat{P}_4$; see below.

Variance estimation

In the Supporting Information, Givens *et al.* (2016) derived the asymptotic distributional properties of $\hat{N}_{2011}$ as extensions to the results of Steinhorst and Samuel (1989), with variance estimates as an extension of Wong (1996); see also Fieberg (2012). For the 2019 abundance estimator, the 2011 derivations can be simplified to derive parallel results, including an asymptotically unbiased variance estimate. The 2019 derivations follow straightforwardly from the previous work, so we present only the final results here.

Define $\theta_i = 1/p_i$. The detection probabilities are estimated in a later section. We use the logit link relationship in the detection probability model, the asymptotic normality of the corresponding parameter estimates $\hat{\boldsymbol{\beta}}$, and the approximation that the corresponding estimated variance-covariance matrix $\hat{\boldsymbol{\Phi}}$ can be treated as known for large samples. If $\hat{\boldsymbol{\beta}}$ and $\hat{\boldsymbol{\Phi}}$ are consistent, then so are the estimators given below. To simplify notation, we define the following terms related to the detection probability model:

$$\hat{\mu}_i = \mathbf{X}_i^T\hat{\boldsymbol{\beta}} \quad \phi_i = \mathbf{X}_i^T\hat{\boldsymbol{\Phi}}\mathbf{X}_i \quad \phi_i = \mathbf{X}_i^T\hat{\boldsymbol{\Phi}}\mathbf{X}_j \quad \tilde{\phi}_{ij} = \phi_i/2 + \phi_j/2 + \phi_{ij}$$

Then

$$\hat{\theta}_i = 1 + \exp\{-\hat{\mu}_i - \phi_i/2\} \tag{1.3}$$

is an asymptotically unbiased estimator for θ_i. Furthermore, it can be shown that

$$\widehat{\mathrm{var}}\{\hat{\theta}_i\} = \exp\{-2\hat{\mu}_i - 2\phi_i\}(\exp\{\phi_i\} - 1) \tag{1.4}$$

and

$$\widehat{\mathrm{cov}}\{\hat{\theta}_i, \hat{\theta}_j\} = \exp\{-\hat{\mu}_i - \hat{\mu}_j - \tilde{\phi}_{ij}\}(\exp\{\phi_{ij}\} - 1) \tag{1.5}$$

are asymptotically unbiased variance and covariance estimators.

There were 61 sightings made with only binoculars for which no sighting distance was obtained (and hence no p_i can be estimated). The remaining sightings included distances determined by theodolite. Following Givens *et al.* (2016), we make the (overly) conservative assumption that detection probability equals 1 in the binoculars-only cases. Then, partitioning and re-indexing the sightings appropriately, we can write

$$\hat{N} = \frac{1}{\hat{E}\hat{P}_4}\sum_{t=1}^{T} c_t\theta_t + \frac{1}{\hat{E}\hat{P}_4}\sum_{b=1}^{B} c_b = \frac{1}{\hat{E}\hat{P}_4}\tilde{N}_t + \frac{1}{\hat{E}\hat{P}_4}\tilde{N}_b = \hat{N}_t + \hat{N}_b \tag{1.6}$$

where *T* and *B* are the numbers of theodolite and binocular sightings, respectively. This leads us to conclude that

$$\widehat{\mathrm{var}}\{\hat{N}\} = \widehat{\mathrm{var}}\{\hat{N}_t\} + \tilde{N}_b^2\,\widehat{\mathrm{var}}\left\{\frac{1}{\hat{E}\hat{P}_4}\right\} + 2\tilde{N}_b\tilde{N}_t\,\widehat{\mathrm{var}}\left\{\frac{1}{\hat{E}\hat{P}_4}\right\} \tag{1.7}$$

where the final term in this expression is a covariance term resulting from the fact that $\mathrm{cov}\{uR, vSR\} = uv(\mathrm{E}S)\mathrm{var}\{R\}$. To derive $\widehat{\mathrm{var}}\left\{\frac{1}{\hat{E}\hat{P}_4}\right\}$ we use the fact that (when *R* and *S* are independent) $\mathrm{var}\{RS\} = (\mathrm{E}R)^2\mathrm{var}\{S\} + (\mathrm{E}S)^2\,\mathrm{var}\{R\} + \mathrm{var}\{R\}\mathrm{var}\{S\}$. We also use the bootstrap estimate for the variance of $1/\hat{E}$ (see a later section) and the approximation that $\widehat{\mathrm{var}}\{1/\hat{P}_4\} = \frac{\widehat{\mathrm{var}}\{\hat{P}_4\}}{(\mathrm{E}\hat{P}_4)^4}$ where $\widehat{\mathrm{var}}\{\hat{P}_4\}$ is estimated in a later section.

What remains, then, is to derive an expression for $\widehat{\mathrm{var}}\{\hat{N}_t\}$. First we note that

$$\begin{aligned}\widehat{\mathrm{var}}\{\hat{N}_t\} = {} & \frac{1}{\hat{E}^2\hat{P}_4^2}\widehat{\mathrm{var}}\{\tilde{N}_t\} + \frac{\tilde{N}_t^2}{\hat{E}^2}\widehat{\mathrm{var}}\{1/\hat{P}_4\} + \frac{\tilde{N}_t^2}{\hat{P}_4^2}\widehat{\mathrm{var}}\{1/\hat{E}\} + \frac{1}{\hat{E}^2}\widehat{\mathrm{var}}\{\tilde{N}_t\}\widehat{\mathrm{var}}\{1/\hat{P}_4\} + \\ & \frac{1}{\hat{P}_4^2}\widehat{\mathrm{var}}\{1/\hat{E}\}\widehat{\mathrm{var}}\{\tilde{N}_t\} + \tilde{N}_t^2\,\widehat{\mathrm{var}}\{1/\hat{E}\}\widehat{\mathrm{var}}\{1/\hat{P}_4\} + \widehat{\mathrm{var}}\{1/\hat{E}\}\widehat{\mathrm{var}}\{1/\hat{P}_4\}\widehat{\mathrm{var}}\{\tilde{N}_t\}\end{aligned} \tag{1.8}$$

and the only term in this expression that we have not explained how to estimate is $\widehat{\mathrm{var}}\{\tilde{N}_t\}$. For this, we return to the Supporting Information of Givens *et al.* (2016) to obtain

$$\widehat{\mathrm{var}}\{\tilde{N}_t\} = \hat{V}_1 + \hat{V}_2 \tag{1.9}$$

where

$$\hat{V}_1 = \sum_{t=1}^{T} c_t^2 (\hat{\theta}_t^2 - \hat{\theta}_t - \widehat{\mathrm{var}}\{\hat{\theta}_t\})$$
$$\hat{V}_2 = \sum_{t=1}^{T} c_t^2 \widehat{\mathrm{var}}\{\hat{\theta}_t\} + \sum_{t \neq s}^{T}\sum c_t c_s \widehat{\mathrm{cov}}\{\hat{\theta}_t, \hat{\theta}_s\} \tag{1.10}$$

Wong (1996) has shown for this type of problem that better confidence intervals are obtained by applying a normal approximation to log abundance and then back-transforming the result. In our case, defining $\widehat{\mathrm{CV}}^2 = \widehat{\mathrm{var}}\{\hat{N}\} / \hat{N}^2$ we can estimate a 95% confidence interval for *N* using

$$(\hat{N}\exp\{-1.96\widehat{\mathrm{CV}}\},\ \hat{N}\exp\{1.96\widehat{\mathrm{CV}}\}) \tag{1.11}$$

Passage estimation, $\hat{E}$

Observer effort was not continuous during the season. At some times the perch was not staffed, usually because environmental conditions were deemed unsafe, the lead was closed and sea ice completely covered the survey area, or visibility was poor or unacceptable (Fig. 2). Let $H_s = 874.1$ the total number of hours during the survey season (from hour 2,483.9 to 3,358.0) and let H_w denote the total number of those hours for which observer watch effort was maintained during qualifying conditions. Since $H_w < H_s$, our abundance estimation must correct for periods of missed survey effort. This correction does not adjust for whales that migrated past the perch before or after the survey season.

The passage rate of whales during the season is not constant. Therefore, correcting for missed effort requires more than simply adding up missed clock time. Recall from equation (1.1) that $\hat{N}$ involves a sum of terms $\hat{h}_i = c_i/\hat{p}_i$ which we call Horvitz-Thompson contributions because $\hat{h}_i$ represents the number of whales that the *i*th sighting contributes to the overall abundance estimate (before correcting for effort and availability). We apply the approach of Givens *et al.* (2016), using the $\hat{h}_i$ to develop a passage rate correction; for completeness we repeat their description below. This approach relies on the fact that passage rate is not correlated with observer presence, as shown from years of experience on the ice, from acoustic and aerial observations during the surveys, and from the traditional knowledge of native hunters in the region.

Let $f_r(t)$ denote the passage rate of whales past the perch (and within 4 km), so that the total number of whales passing within this range – seen or unseen – between times t_1 and t_2 is $\int_{t_1}^{t_2} f_r(t)\,dt$. Let *S* and *W* denote the sets of time periods corresponding to the total analysed survey season and periods of qualifying watch, respectively. Then the proportion of passage during qualifying watch is

$$E = \int_W f_r(t)\,dt \Big/ \int_S f_r(t)\,dt \tag{1.12}$$

and by estimating this quantity we can derive the desired correction factor $1/\hat{E}$.

To estimate the correction factor, we bin the $\hat{h}_i$ into 12 hour time blocks, $B_1,\ldots,B_{77}$, and define $\hat{H}_j$ to equal the sum of all the $\hat{h}_i$ observed during block B_j. Thus $\hat{H}_j$ is the total Horvitz-Thompson contribution for the *j*th block, i.e. an estimate of the (uncorrected) total number of whales passing with 4 km of the perch during periods of qualifying effort during that block. Let T_j denote the total amount of qualifying watch effort during block *j*, and let the blocks be referenced by their temporal midpoints t_j. Define $\hat{R}_j = \hat{H}_j/T_j$ to be the number of passing whales (within 4 km) per qualifying watch hour in the *j*th block.

We model the block passage rates using a shifted gamma generalised additive model (GAM) with log link, specifically $\hat{R}_j + s \sim \mathrm{gamma}(\gamma_j,\eta_j)$ where the mean of $\hat{R}_j + s$ is $\xi_j = \gamma_j,\eta_j$ and

$$\log \xi_j = f_r^*(t_j) \tag{1.13}$$

where $f_r(t) = \exp\{f_r^*(t_j)\}$ is a smooth rate function. The shift constant *s* was necessary to cope with the fact that some $\hat{R}_j = 0$. The value of *s* was chosen to maximise the explained deviance.

The model was fit using the mgcv package in the **R** computing language (Wood 2004, 2011, 2017; R Core Team, 2019). We used $k = 20$ knots, which allows for good fidelity to the data without overfitting. The default generalised cross-validation method was used to choose the smoothness penalty. Points were weighted proportionally to the T_j.

This model can be re-expressed in terms of a matrix $\mathbf{U}$ with one row per spline basis function and the *j*th column representing the *j*th block, using $\log \xi_j = \mathbf{U}^T\gamma$. The column vector of parameter estimates $\hat{\gamma}$ has a limiting distribution $\hat{\gamma} \sim N(\gamma, \Lambda)$. Technically, this is a limiting Bayesian posterior distribution, but no prior information about γ or the $\hat{R}_j$ is incorporated in the analysis beyond the smoothness penalty; see Wood (2017). An estimated covariance matrix $\hat{\Lambda}$ is obtained while fitting this GAM.

The final step is to estimate *E*. Define

$$\hat{E} = \int_W \hat{f}_r(t)\,dt \Big/ \int_S \hat{f}_r(t)\,dt \tag{1.14}$$

where the integrals are approximated using Simpson's rule (e.g. Givens and Hoeting, 2013).

We derived $\widehat{\mathrm{var}}\{1/\hat{E}\}$ using the parametric bootstrap approach recommended by Wood (2006, pp. 202–203). The GAM is first fit to the original data. Using the estimated mean function from this fit, bootstrap response data are generated from the parametric gamma model. A new GAM is fit to these data to obtain a bootstrap estimate of the smoothing parameter. Next, a GAM is fit to the original data using the bootstrap smoothing parameter value. This produces one set of bootstrap pseudo-estimates $\hat{\gamma}^*$ and $\hat{\Lambda}^*$. We performed 2,500 such bootstrap iterations. Then, to simulate the bootstrap distribution of $1/\hat{E}$, we select a random one of the 2,500 distributions $N(\hat{\gamma}^*, \hat{\Lambda}^*)$ and sampled a value $\hat{\gamma}^*$ from it. This value was used to obtain a bootstrap pseudo-value $\hat{E}^*$ using equation (1.14). From the set of $\hat{E}^*$ generated in this way we compute the sample variance of the set of values of $1/\hat{E}^*$ to derive $\widehat{\mathrm{var}}\{1/\hat{E}\}$.

Although $1/\hat{E}$ and its variance estimator are not statistically independent of the $\hat{p}_i$, we believe that the nature of $\hat{E}$ as an integral of a smooth function of the huge set of those $\hat{p}_i$ provides sufficient justification to treat $1/\hat{E}$ as approximately independent for the purposes of deriving the variance of $\hat{N}$ (equation 1.8).

Detection estimation, p_i

We used the published model and results by Givens *et al.* (2015) for the 2011 survey without modification. The 2011 survey included a comprehensive two-perch independent observer experiment that provided detailed data for estimating detection probabilities. Fig. 4 shows the results of that analysis (Givens *et al.*, 2015). Detection probability was found to depend on sighting distance, lead condition and group size. In this model, parameter estimates $\hat{\boldsymbol{\beta}}$ describe how detection probabilities depend on covariates. The corresponding estimated variance-covariance matrix for these parameter estimates is denoted $\hat{\boldsymbol{\Phi}}$, and can be derived from the model fit.

Availability estimation, $\hat{P}_4$

Fig. 3 shows the data used to estimate availability. We fit an inverse variance weighted linear regression to these data and calculated $\hat{P}_4$ as the point on the fitted line (shown) at 1.29nm north of Point Barrow, the latitude of the 2019 perch. There are two potential standard errors for this estimate: for the estimation of the mean and for the prediction of a new case (Kutner *et al.*, 2005). The latter is much larger because it incorporates the variance for the mean and the variance of a new case around the fitted mean. It is somewhat of a default to use the standard error of the mean, as would be equally tempting if one were simply averaging some past values for $\hat{P}_4$. However, we believe this approach is incorrect. We believe there is significant interannual variation in P_4 *even if perch location was unchanged*, and therefore a proper expression for the uncertainty in $\hat{P}_4$ is the standard error of prediction. We denote the square of this estimate as $\widehat{\mathrm{var}}\{\hat{P}_4\}$. Zeh and Punt (2005) show interannual change in P_4 with changing perch locations, and Fig. 4 of Givens *et al*. (2016) shows intra-annual change. The magnitude of both types of variation is large.

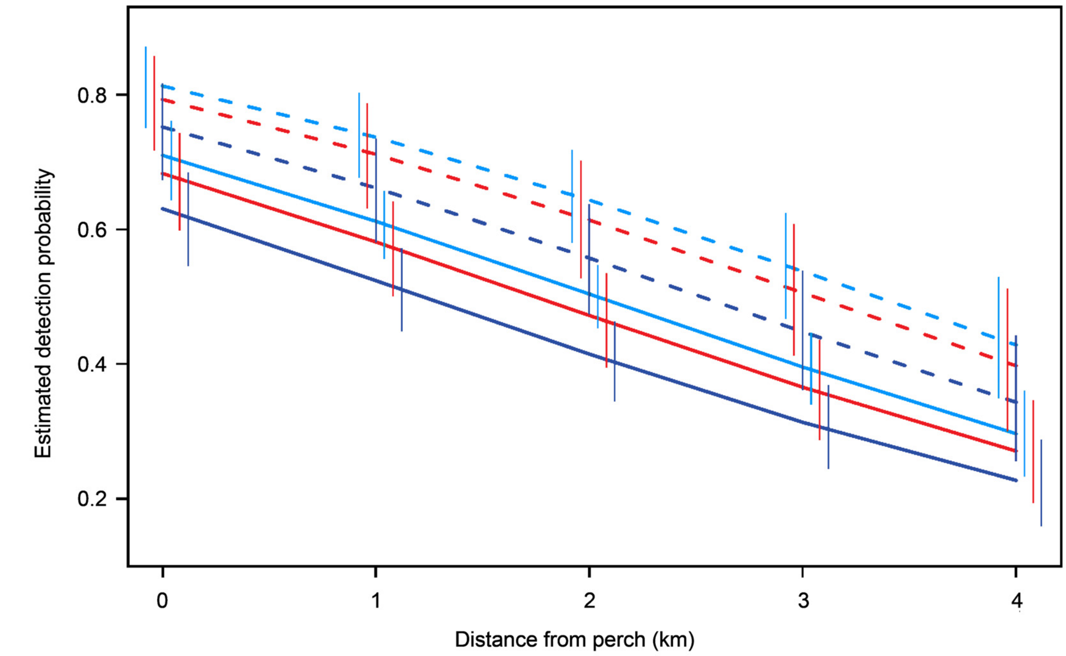

Fig. 4. Estimated detection probabilities from Givens *et al.* (2015). Solid lines correspond to sightings of single animals; dotted lines are for groups. For each line style, there are 3 colors: continuous leads (light blue), patchy leads (red), and wide open water (dark blue). Individual 95% confidence intervals for the estimates are shown as vertical lines at integer km distances, but horizontally offset (and slightly vertically shifted to compensate) so they can be distinguished.

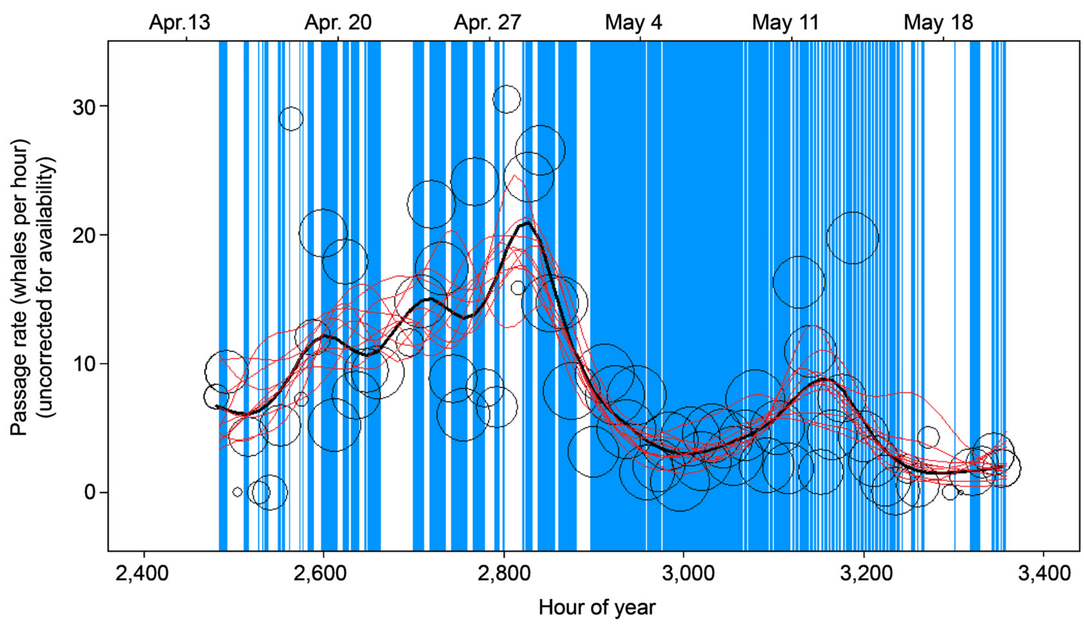

Fig. 5. Estimated passage rates, $\hat{R}_j$, for 12h blocks during the survey season. The area of the circle is proportional to the qualifying watch time, T_j, and the weight each point receives in the GAM model. The fitted GAM is shown with the black curve, and the red curves are 10 bootstrap replicates. The blue bands indicate periods of qualifying watch.

RESULTS

Fig. 5 shows the estimated passage rates for 12h blocks during the survey season. The black line is the estimated passage rate function. Ten bootstrap replicates are shown in red. We estimate that 59.9% of whales passing within 4 km of the perch during the survey season did so during periods of qualifying watch ($\hat{E} = 0.599$). This leads to a correction factor of $1/\hat{E} = 1.67$ with $\widehat{\text{var}}\{1/\hat{E}\} = 0.0337^2$.

The regression of availability on distance north of Point Barrow (Fig. 3) is highly statistically significant ($R^2 = 0.87$, $F_{1,6} = 39.68$, $p = 0.0007$). The estimated availability correction for 2019 is $\hat{P}_4 = 0.5713$. The standard error of prediction for P_4 is $(\widehat{\mathrm{var}}\{\hat{P}_4\})^{1/2} = 0.1263$. For comparison, if we had used the standard error of the mean, the value would be 0.0529.

The point estimate for 2019 abundance is $\hat{N} = 12{,}505$. Variance estimates were as follows. For theodolite data, $\hat{V}_1 = 75.672^2$, $\hat{V}_2^2 = 14.162^2$, $\widehat{\mathrm{var}}\{\tilde{N}_t\} = 171.73^2$, and, after incorporating the variance components for $1/\hat{P}_4$ and $1/\hat{E}$, $\widehat{\mathrm{var}}\{\hat{N}_t\} = 2{,}823.56^2$. After adjusting for binocular data, the final estimates are $\widehat{\mathrm{var}}\{\hat{N}\} = 2{,}854.17^2$ and $\widehat{\mathrm{CV}} = 0.228$. The 95% confidence interval for abundance is (7,994, 19,560).

DISCUSSION

The 2019 survey was unusual in the sense that the migration's timing and pattern were considerably different than has been observed in recent surveys and vastly different than the early years of the ice-based surveys in the 1980s and 1990s.

The most notable characteristic of the 2019 bowhead whale abundance estimate of 12,505 (7,994, 19,560) animals is that it is substantially lower than the 2011 estimate of 16,820 (15,176, 18,643) animals (Givens *et al.,* 2016). The 2019 estimate is also considerably less precise than the 2011 estimate. However, we do not believe that the 2019 result represents an actual, substantial reduced abundance of bowhead whales for several reasons:

(1) The 2019 confidence interval fully encloses the 2011 interval, so our new result is consistent with a hypothesis of 'no change' (or even 'increased abundance');

(2) The bowhead migration has occurred increasingly early in recent years (George *et al.*, 2013). In 2019, we began the survey with brief watches on 16, 17 and 23 March. However, leads closed in late March and only reopened on 14 April. Based on the times that bowheads passed other villages south of Utqiaġvik, the detection of bowhead calls on a recording device near the perch on 7 April, and the fact (Figs 2 and 5) that bowheads were numerous as soon as watch began on 14 April rather than trickling in at the start, we believe that substantial numbers of bowheads passed the perch location before watch (re)opened on 14 April. Whales passing before then (and after watch ended on 23 May) are not included in our abundance estimate. The total hours of watch conducted in 2019 were less than 75% of what was achieved in 2011;

(3) During some periods of high passage, the bowheads seemed to migrate along routes that tended to be far from the perch. There were some huge, perhaps unprecedented, shore leads (nearly 100 km in width). To the southwest the shore-fast ice broke away, and bowheads tended to follow a straight line (from Pt. Franklin to Pt. Barrow) along the distant pack ice rather than a curved path that would keep them near the shore-fast ice and hence near the perch; see Fig. 6. Some senior whale hunters feel the lack of ice between Pt. Franklin and Pt. Barrow (Fig. 6) and the narrow shore-fast ice shelf was one of the biggest factors influencing the unusual whale distribution, with whales being farther offshore compared to previous years. If more whales were further offshore than would be 'typical' for the 2019 perch location, then our estimate of P_4 would be biased upward, resulting in downward bias in estimated total abundance;

(4) The survey was hindered by substantial interference from power boats. In most previous survey years, hunters quietly paddled traditional umiaqs (i.e. seal skin covered boats) for much of spring whaling. However, in 2019, hunters consistently used power boats during the spring whaling season, in all regions of the coast near Utqiaġvik, including a camp just 200 metres north from the perch. The disturbance to the survey was especially severe starting 7 May to the end of the season. when the entire coast was opened to powered skiffs. The increased use of power boats was necessary for the community to land whales because of the unusual distribution of bowheads. The whales were inaccessible to skin boat hunters at the shore-fast lead edge, and much longer distances needed to be covered to hunt bowheads

(and tow them back to the shore-fast ice). Bowheads avoid power boats by diverting further offshore and/or reducing surface times (Richardson and Malme, 1993). This behaviour persists for a period of time after the disturbance. Perch observers reported a downturn in the whale count in the hours after power boats launched. Givens *et al.* (2021) estimate a correction factor for the abundance estimate presented here, to correct for boat noise disturbance. Their corrected abundance estimate was about 12% higher than the result in this paper;

(5) The 2019 perch was one of the shortest primary perches in the last 40 years. The perch height was 6.2m, compared to an average of 8.2m. Only for part of the season in 1981 has a shorter primary perch ever been used. This means that our detection probability estimates ($\hat{p}_i$) may be too high since they are based on independent observer sight-resight data from the higher 2011 perch (10.5m). Furthermore, the estimated proportion of whales passing within visual range of the 2019 perch ($\hat{P}_4$) may be too high if the effective visual range was actually diminished due to the low perch height. Both these potential biases would cause downward bias in the abundance estimate;

(6) Huntington *et al.* (2020) offer a comprehensive description of the perturbations of Bering and Chukchi Sea marine environment in 2017–2019. The changes include unexpected deviations to the bowhead's habitat, distribution, and behaviour, based on comparative data from satellite telemetry and other sources. The abrupt changes to the marine ecosystem and record low sea-ice cover that occurred during this period were not predicted. The potential impact of these changes on our survey is unknown;

(7) Bowhead health assessment information for 2019 suggest that harvested bowheads did not exhibit obvious reductions in health condition, and the ice-based and summer/autumn aerial surveys indicated high calf production (Stimmelmayr *et al.*, 2020). There is no evidence of an unusual mortality event; and

(8) Our treatment of whales sighted only with binoculars (i.e. assuming that detection probability equals 1 for such cases) introduces downward bias in the abundance estimate.

Unlike 2011, there were no long periods of heavy fog in 2019. Also, east winds were nearly constant in 2019 and hence the leads seldom closed. In these respects, the 2019 survey was more 'successful' than many previous ones in terms of consistent open leads in good weather. However, if whales are not available for visual detection from the perch for the reasons given above, this cannot be remedied by having open leads and good visibility.

As mentioned previously, we did not estimate P_4 from acoustic data in 2019. Our results show that a sufficiently precise estimate could be obtained from a regression of $\hat{P}_4$ values from past years on perch location. However, if the relationship shown in Fig. 3 is spurious, or if 2019 did not follow that pattern for some reason (such as the anomalous ice/lead conditions described above), then the mis-estimation of P_4 could be substantial, perhaps leading to serious under-estimation of total abundance. We attempted to include ocean depth at the perch site as a covariate in this regression, but there was not a strong or statistically significant relationship. The

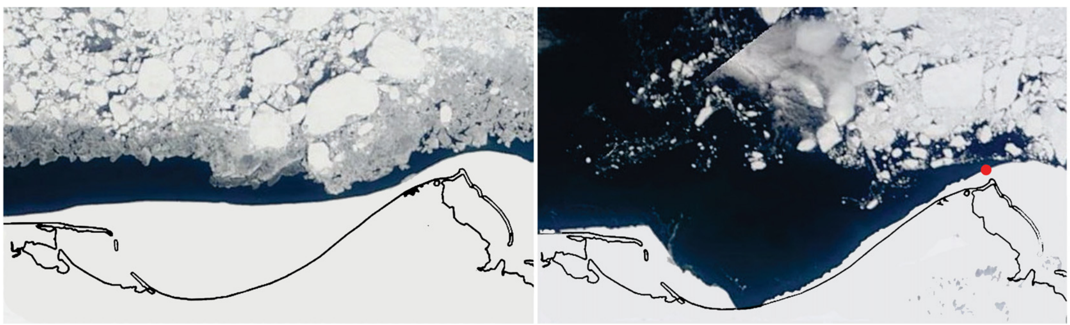

Fig. 6. Satellite images of the region near Pt. Barrow on 13 May 2011 (left) and 2019 (right). In 2019, hunters reported that bowheads sometimes tended to migrate through the open water far from the shore-fast ice edge (and the perch). The 2019 perch location is shown with the red dot in the right image. Some local whale hunters believe the shore-fast ice configuration explained much of the differences in whale distribution in 2019, and why bowheads swam further offshore. The horizontal span of these images is about 150 km.

configuration of the ice, lead, and coastline appears to be more important. This suggests that bowheads may begin to continue swimming in a more northerly course near Point Barrow as the shore-fast ice-shelf bends east. Note that we chose the conservative approach of using the standard error of prediction (not estimation) from this analysis, which made the magnitude of the standard error 2.5 times larger for this correction factor compared to the alternative. We believe that this more correctly represents the inherent uncertainty about P_4.

Although the data point for 2011 in Fig. 3 is visually somewhat anomalous, it does not unduly affect the estimate of P_4. The Cook's distance value for this point is 0.039, corresponding to the 0.038 percentile of a $F_{2,6}$ distribution. Kutner *et al.* (2005) suggest that a corresponding percentile of less than 0.10 or 0.20 implies "little apparent influence" (p. 403), while a percentile exceeding 0.50 implies substantial influence. Additional regression diagnostics not reported here also confirm that the regression is reliable and deleting 2011 entirely from the regression would have changed the estimate of P_4 from 0.571 to 0.578. The regression approach is an improvement from previous practice, where a simple average of past P_4 values would be used when no contemporaneous direct estimate was available. Note that the regression prediction is, algebraically, a weighted average of past values.

Our experience in 2019 raises the question of whether the ice-based survey protocol remains viable in the changing Arctic. Sea ice is getting thinner, less predictable, and thus more difficult and less safe for observers. Typically, more bowheads are detected on these surveys than by any other means, including aerial surveys. Deteriorating ice conditions require us to terminate the survey sooner than we would prefer, or than we would have in past decades. Surveys in the 1980s typically continued into June. However, the migration is tending to be earlier each year, and despite limitations due to more abundant ice early in the season, poorer weather, and scarce sunlight, we have attempted to move the survey earlier too. If 'everything goes right', we believe that a successful ice-based survey abundance estimate coupled with acoustic monitoring is likely to be more precise than any other approach, provided that a precise estimate of P_4 is made.

Other survey methods are being contemplated if continued monitoring of the bowhead population from an ice-based survey is not possible in the future. In 2019, an aerial survey was flown across the Beaufort Sea from Point Barrow to Amundsen Gulf. This survey was supported by the U.S. National Marine Fisheries Service, the U.S. Bureau of Ocean Energy Management and the North Slope Borough. The survey was successful and the results were reported to the IWC Scientific Committee by Ferguson (2020). In the future, remote sensing via satellite imagery (Cubaynes *et al.*, 2020) might be considered, too.

Clearly 2019 was not a perfect year, and the abundance estimate reflects this. However, we believe that the estimate is sufficiently accurate, notwithstanding its limitations, and precise for use with the IWC's Bowhead Strike Limit Algorithm, the computational tool the IWC Scientific Committee uses to provide advice on hunting quotas (IWC, 2003a). That procedure has been shown to be extremely robust to the sorts of biases and uncertainties discussed here (IWC, 2003b). Furthermore, the precision of our estimate is adequate: the Bowhead Strike Limit Algorithm was tested with abundance estimate CVs ranging from 0.10 to 0.34, with 0.25 being the baseline case (IWC, 2003b). Our $\widehat{\mathrm{CV}} = 0.228$ fits nicely inside this tested parameter space.

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