Generalized linear models applied to stranding data of South American sea lions (*Otaria flavescens*) and South American fur seals (*Arctocephalus australis*) in southern Brazil [1]

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Abstract: Monthly strandings of South American sea lions, *Otaria flavescens,* and South American fur seals, *Arctocephalus australis,* collected between 1993 and 2002 along the coastline of Rio Grande do Sul, Brazil, are analysed with generalized linear models (GLM) to describe spatial and temporal stranding patterns and to estimate the annual number of stranded animals by species. Results indicate a strong seasonal pattern for both species with maximum average number of strandings in September and minimum in January. While year effects are not significant, differences between regions (north or south of the Patos Lagoon mouth) are. Estimated numbers of stranded sea lions and fur seals in a typical year are 115, with 95% confidence interval (76; 160) and 95, with 95% confidence interval (48; 180), respectively. GLMs proved to be a very useful tool to examine the influence of time and space on strandings in a statistically rigorous framework.

Resumo: Encalhes mensais de leões marinhos do sul, *Otaria flavescens,* e lobos marinhos do sul, *Arctocephalus australis,* registrados do ano de 1993 até 2002 ao longo da costa do Rio Grande do Sul, Brasil, são analisados com modelos lineares generalizados (GLM) a fim de descrever padrões espaciais e temporais bem como estimar o número anual de encalhes efetivos por espécie. Os resultados indicam um forte padrão sazonal para ambas as espécies com máximo de encalhes em setembro e mínimo em janeiro. Enquanto efeitos anuais não são significativos, diferenças regionais (ao norte e ao sul dos molhes da Lagoa dos Patos) o são. Os números anuais de encalhes estimados para leões e lobos marinhos são 115 e 95, respectivamente, com intervalo de confiança de 95% (76; 160) e (48; 180). GLMs se mostraram ferramentas úteis para examinar de forma estatisticamente rigorosa, as influências de tempo e espaço nos encalhes de pinípedes.

Keywords: generalized linear models, pinniped, strandings, South American sea lion, South American fur seal, *Otaria flavescens*, *Arctocephalus australis*, southern Brazil.

Introduction

Ordinary linear regression has played a central role in classical applied statistics due to the possibility to model empirically some stochastic response variable (Y) as a linear function of a set of explanatory variables ($x_1, x_2,..., x_k$). However, as the range of applications grew with time, restrictions imposed by the requirements of a Gaussian probability distribution for Y and constant variance over the range of observed values of ($x_1, x_2,..., x_k$) became apparent. Generalized linear models (GLMs; McCullagh and Nelder, 1989; Dobson, 2002) attempt to accommodate both restrictions simultaneously by extending linear regression in an effective way.

The components of a GLM are as follows:

(i) the response variable Y follows a probability distribution of known form within the exponential family of distributions. This family includes continuous distributions like the Gaussian and gamma, as well as discrete distributions like Poisson, binomial and negative binomial;

(ii) the distribution of Y depends on a set of explanatory variables through a single linear predictor, $\eta = \sum_{j=1}^{k} x_j \beta_j$, where x_j are known functions of a selection of k explanatory variables and β_j unknown parameters to be estimated from the data;

(iii) the linear predictor η is related to the mean of Y, $\mathrm{E}(Y) = \mu$ by a known *link function g*; namely $\eta = g(\mu)$;

(iv) the variance of Y is a function of μ and a scale parameter ϕ. For the Poisson distribution $\phi = 1$, while for the negative binomial or the Gaussian distributions, ϕ is usually unknown and has to be estimated from the data.

In light of the above definitions, it is clear that linear regression is just a special case of GLMs that results when the distribution of Y is Gaussian and the link function is identity $\eta = \mu$. In contrast, stranding data are counts (0, 1, 2, 3,…) for which discrete probability distributions such as the Poisson or negative binomial are better candidates. The link function in both cases is the natural logarithm, namely $\eta = log(\mu)$.

For an excellent overview of GLMs and other related modeling processes the reader is referred to Venables and Dichmont (2004). Additional theoretical considerations can be found in Dobson (2002) and Ripley (2002).

An up-to-date review of published accounts on pinniped occurrences along the Brazilian coastline over the last 30 years is given in Silva (2004). Despite its historical value, this collection of isolated events neither allows for comparative studies in time or space, nor quantifies the number of strandings in a statistically rigorous way. There are studies about the dynamics of movements

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[1] The term "stranding" is generally used when the animal(s) is (are) still alive when reaching the beach, however, here the term will also refer to dead animals that were washed ashore.

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along the Rio Grande do Sul coast, and occupancy patterns of wintering concentrations at the Patos Lagoon's east jetty (Molhe Leste) and the Lobos island (Ilha dos Lobos) for sea lions *Otaria flavescens* by Rosas (1989), Rosas *et al.*(1994) and in several unpublished technical reports. Due to its coastal distribution and feeding behavior, *O. flavescens* suffers aggression from fishers and aquaculture workers throughout its range (Silva, 2004). Along the Rio Grande do Sul coast the interaction with fisheries is pointed out as the main cause mortality in this species (Rosas, 1989; Pinedo, 1990; Silva, 2004).

This article illustrates the use of GLMs to describe the observed stranding of South American fur seals (*Arctocephalus australis*) and South American sea lions (*Otaria flavescens*) along the coastline of Rio Grande do Sul, southern Brazil, according to location, time and search effort (*i.e.* percentage coastline coverage). The analysis is aimed at (*i*) describing the spatial and temporal patterns in strandings and (*ii*) estimating the absolute number of annual stranding for both species.

Methods

Study Area and Data Collection

The Rio Grande do Sul coast is a 623km long sandy shoreline between Torres (29°21'S) and Chui (33°45'S), south Brazil (Figure 1). The region is under direct influence of the subtropical convergence of the western South Atlantic Ocean and is considered one of the most productive fishery grounds in Brazil (Seeliger *et al.*,1997).

Between January 1993 and December 2002 a total of 182 expeditions were conducted along the coast. Of these, 98 expeditions covered the region to the south of the Patos Lagoon jetties and 84 covered the northern region with an average search effort of 2746.2km/year (sd = 496km). The percentage coverage per month is described in Table 1.
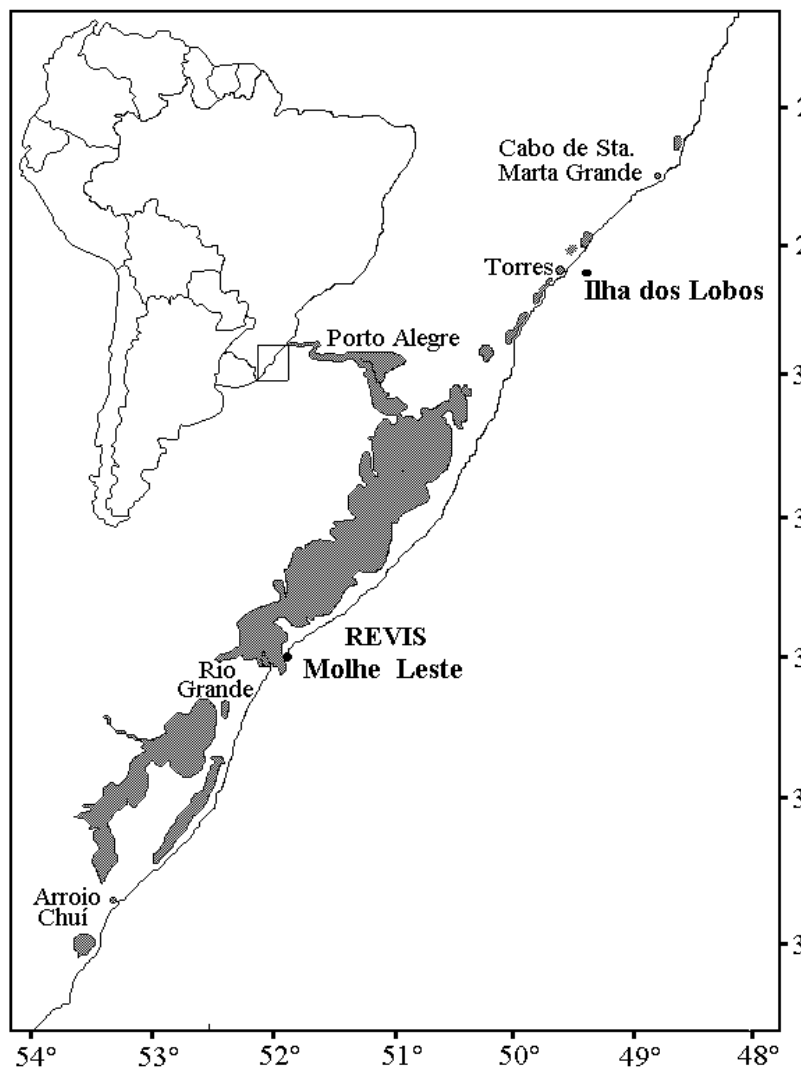

Figure 1. The study area: the coastline of Rio Grande do Sul, Brazil, between Torres and Arroio Chuí highlighting the Molhe Leste wildlife reserve (REVIS = Reserva da Vida Silvestre).

Table 1. Percentage coverage of coastline by month and year (1993-2002).

MONTH	YEAR									
	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002
Jan	35.3	27.1	20.9	61.6	3.9	54.9	*	18.0	*	57.3
Feb	60.2	21.7	10.6	35.3	62.1	22.6	*	17.8	*	57.0
Mar	*	35.3	25.4	74.3	35.3	100.0	22.0	17.0	*	76.7
Apr	12.8	10.6	11.6	36.1	35.3	36.3	8.0	17.7	100.0	*
May	25.8	10.6	61.8	58.1	39.3	43.3	37.1	19.3	100.0	57.5
Jun	22.0	*	68.7	57.8	35.3	22.2	1.9	17.3	100.0	17.7
Jul	34.3	35.3	43.2	1.9	47.8	80.4	28.6	19.1	100.0	57.0
Aug	20.7	9.6	40.6	71.7	41.7	50.9	17.2	1.9	45.6	1.9
Sep	1.9	36.9	90.4	50.7	60.7	61.2	31.8	1.9	42.4	68.1
Oct	35.3	*	67.3	36.1	71.4	23.6	23.0	5.3	100.0	*
Nov	5.3	57.9	60.5	64.2	49.0	48.3	20.5	1.9	100.0	93.7
Dec	5.3	46.9	29.1	55.5	62.0	16.1	23.0	3.9	80.9	18.5
Avg[(1)]	23.5	29.2	44.2	50.3	45.3	46.7	22.8	11.8	85.4	50.5

(*) month without expedition; (Avg[(1)]) average monthly coverage.

Expeditions with two or three observers were carried out on a four-wheel drive vehicle travelling at an average speed of about 45km/h. Along with the counting of all stranded pinnipeds by species, gender of each specimen was determined based on external characteristics, and biometric measurements were also taken. Location along the coastline and evidence of interaction with fisheries were also registered. Each sampled carcass was marked with spray to avoid double counting in later expeditions. Species identification followed the regional field guide by Pinedo *et al.* (1992). All dead animals were classified according to their stage of decomposition in a scale from 1 to 5 as follows: (G1) recent death; (G2) rigor mortis and intact carcass; (G3) carcass with strong odor and fluid releases; (G4) putrifying and slightly destroyed carcass; (G5) strongly putrified and destroyed carcass. Detailed descriptions of these data are available in Silva (2004). For the present study, only carcasses that could be identified as *O. flavescens* and *A. australis* were considered.

Statistical Analysis

We assume as our working hypothesis that the number of observed stranded animals (Y) of a given species is possibly affected by the explanatory variables month, year, location along the coast and sampling effort. By modelling Y as a function of these explanatory variables, we expect that the effect of each can be quantified and statistically examined. Furthermore, we plan to use the adjusted model to predict strandings for those months in the time series which do not have any observations or that have only partial coverage of the coastline. Because *Y*s are counts, hence discrete random variables, the use of a GLM seems appropriate. Two candidate probability distributions are examined.

$$Y_i \sim \text{Poisson}\ (\mu_{i,}\ \phi = 1)$$
$$Y_i \sim \text{Negative binomial}\ (\mu_{i,}\ \phi)$$

for all data points $i = 1, \ldots, n$.

Furthermore, we assume a linear relation between $\eta_i = log\ (\mu_i)$ and the explanatory variables months (1 to 12), year (1993 to 2002), region (south and north of the Patos Lagoon jetties) and effort (kilometers of coastline covered). It should be noted that variables month, year and region are categorical (with 12, 10 and 2 levels, respectively) while effort is quantitative.

Five alternative models are proposed:
Model 1: η = constant + month + region + effort
Model 2: η = constant + month + region + region x effort
Model 3: η = constant + month + region + year + effort
Model 4: η = constant + month + region + year + region x effort
Model 5: η = constant + month + region + region x month + region x effort

Model 1 considers only the effects of month, region and amount of effort on changes in η. Model 2 is an extension of Model 1, by assuming that both regions have different angular coefficients for effort. Model 3 extends Model 1 by including the year effect while Model 4 is a similar extension of Model 2. Finally, Model 5 extends Model 2 with the inclusion of region-month interactions. It should be noted that the number of model parameters (β) grows from Model 1 to Model 5; namely 14, 15, 23, 24 and 26. Model 1 is said to be nested in Model 2 which is nested in both Model 4 and Model 5. Model 1 is also nested in Model 3, whereas Model 2 is not. The advantage of this structured construction of models will become apparent in statistical hypotheses testing described below.

Model selection

The combination of two choices of probability distributions for *Y* with five models for the linear predictor η produces ten possible candidate models. Therefore, a procedure to empirically select a *best model* is necessary. Two information-theoretic criteria are commonly used to assist this choice: Akaike's information criterion (AIC; Burnham and Anderson, 2002) and the Bayesian information criterion (BIC; Schwartz, 1978). To search for a better model is thus equivalent to looking for smaller values of AIC or BIC.

$$BIC = -2\log \hat{L} + p \log n$$

Both criteria use the likelihood function of Ys evaluated at its maximum ($\hat{L}$) and differ in their additional penalty term. While AIC only penalizes the total number of model parameters (*p*), BIC also includes the number of observations (*n*). One consequence of this difference is that, as the number of data points increases, it is more difficult to accept more parameters with BIC than with AIC.

Another important criterion to compare nested models using standard statistical hypotheses testing is the deviance. For a given model M and scale parameter ϕ, the deviance (D_M; Venables and Dichmont, 2004) is calculated in reference to the saturated, most general model S:

$$D_M = 2\left(\log \hat{L}_S (\phi) - \log \hat{L}_M (\phi)\right)$$

A more technical description of deviance is not intended here and can be found in the specialized literature on GLMs (*e.g.* Dobson, 2002, p.77).

Estimating the number of annual strandings

After selecting some "best model", predicted estimates of the average number of strandings $\hat{\mu} = \exp(\hat{\eta}_i)$ for all months within the study period are calculated by assuming a complete coverage of the study area (this is achieved by setting effort such that the coastline is completely covered: *i.e.* effort = 220km in the southern region and effort = 403km in the northern region). This procedure produces estimates of expected strandings which are comparable in time and between regions.

To estimate the absolute number of yearly stranded animals, we sampled *Y*s from 24 negative binomial distributions, each with some parameter $\hat{\mu}$, denoting a particular combination of month and region (12 x 2 = 24 cases), and the scale parameter ϕ, and where both parameters were estimated based on the final selected model. The summation of these 24 simulated *Y*s gives an estimate of the total number of yearly stranded animals. The procedure just described was replicated 1000 times in order to assess variability and obtain bootstrap confidence intervals.

All statistical computations were performed using the software R version 2.0. (R Development Core Team, 2004).

Results

Model selection

Some preliminary residual analysis of the data indicated an outlier observation of 52 stranded animals that occurred in the southern region during October 1993 for *O. flavescens*. Similarly, an outlier observation of 47 stranded animals was identified in the southern region in December of 1997 for *A. australis*. Both were removed from the data set before performing the statistical analysis below.

The values of AIC and BIC for all models are summarized in Table 2. Two features emerge clearly for both species. First, the negative binomial distribution fits consistently better than the Poisson. Second, given a negative binomial distribution for *Y*, the best model based on AIC and BIC are Model 4 and Model 2, respectively. This comes as no surprise since it is known that BIC favors more parsimonious models then does AIC. All analyses that follow refer to model M07 only (*i.e.* Model 2 with a negative binomial distribution for *Y*).

Parameter estimates

The selected GLM fitted to stranding data of *O. flavescens* and *A. australis* produced the parameter estimates listed in Table 3 (columns 2 and 5). The baseline estimate of the logarithm of the expected number of strandings $\eta = \log \mu$ south of the jetties (region 1 (s)) in January (month 1) is given by the intercept. This value is adjusted differently for the amount of effort in both regions as given by the coefficients (effort (s) and effort (n)). Finally, increments for different months of the year (month 2 to month 12) are also added.

For *O. flavescens*, the months of May, June and, most importantly, August until November (with peak in September) have statistically significant increments in η. The effect of each unit of effort induces different increments in each region as indicated by different coefficients. For *A. australis* the monthly effects are similar with significant increments in η between June until November (with peak in September). The coefficient adjusting for amount of effort is only significant in the southern region. For both species the effect for the northern region (region 2 (n)) is statistically unimportant.

The last line of Table 3 displays the maximum likelihood estimates of ϕ and their standard errors. While the estimate 2.361 for *O. flavescens* is in clear support of a negative binomial over-dispersion, because it is more than two standard errors away from $\phi = 1$, the estimate 0.999 for *A. australis* is not.

Table 2. Akaike information criterios (AIC) and Bayesian information criterion (BIC) for models fitted to stranding data of *O. flavescens* and *A. australis*, south Brazil.

MODEL (DISTRIBITUION FOR Y)	O. flavescens		A. australis	
	AIC	BIC	AIC	BIC
M01. Model 1 (Poisson)	834.24	907.10	1008.60	1081.46
M02. Model 2 (Poisson)	785.13	863.19	974.04	1052.10
M03. Model 3 (Poisson)	805.66	925.35	834.37	954.07
M04. Model 4 (Poisson)	764.71	889.60	787.71	912.61
M05. Model 5 (Poisson)	777.68	912.99	906.44	1041.74
M06. Model 1 (neg. binom.)	740.93	813.79	707.63	780.49
M07. Model 2 (neg. binom.)	720.74	**798.80**	701.26	**779.32**
M08. Model 3 (neg. binom.)	742.02	861.72	689.84	809.54
M09. Model 4 (neg. binom.)	**718.52**	843.42	**675.53**	800.43
M10. Model 5 (neg. binom.)	730.40	865.71	694.73	830.03

Obs.: Numbers in bold represent the minimum for each column.

Table 3. Estimates, standard errors and p-values of coefficients for selected generalized linear model (Model 2 with negative binomial distribution) fitted to the number of stranded sea lions (*O. flavescens*) and fur seals (*A. australis*) in south Brazil.

PARAMETER	O. flavescens			A. australis		
	ESTIMATE	STD. ERROR	P-VALUE	ESTIMATE	STD. ERROR	P-VALUE
intercept	-1.42	0.44	.0015	-2.36	0.79	.0032
month 2	-0.65	0.57	.2587	1.49	0.84	.768
month 3	0.44	0.47	.3429	-0.02	0.99	.9878
month 4	0.82	0.47	.0840	1.28	0.87	.1458
month 5	0.86	0.43	.0497	1.50	0.81	.0649
month 6	0.91	0.45	.0466	2.30	0.80	.0047
month 7	0.76	0.43	.0803	3.03	0.77	.0001
month 8	1.45	0.43	.0009	3.24	0.78	.52E-6
month 9	1.82	0.41	.18E-6	3.83	0.77	.15E-7
month 10	1.55	0.44	.0006	3.09	0.79	.0001
month 11	1.65	0.41	.95E-6	2.17	0.78	.0062
month 12	0.63	0.47	.1796	0.45	0.96	.6437
region (n)	0.51	0.33	.1195	0.54	0.44	.2310
effort (s)	0.010	0.0014	.52E-13	0.0073	0.0020	.0003
effort (n)	0.002	0.0010	.0304	0.0005	0.0015	.7571
Ø	2.361	0.546	-	0.9993	0.1731	-

Residual Analysis

Residuals are used to examine the fit of the selected model. The graphical display of residuals against observed strandings (Figure 2; a and c) shows that standardized residuals are within an acceptable range (-3 and 3). Since Y represents discrete counts while the estimates are non-negative expectations, there is no surprise that for small values of Y residuals are mostly negative. Particularly for Y = 0, negative residuals are the only possibility.

The quantile-quantile plots (Figure 2; b and d) further indicate that Gaussian distributions of residuals are acceptable, since the scatterplots are nearly linear.

Simulated Annual Number of Strandings

Standardized estimates of the expected number of strandings $\hat{\mu}$ assuming complete coverage of the shoreline in all months and years (x_4 = 220km in the south (region 1) and x_4 = 403km in the north (region 2)) are shown in Figure 3. These estimates are comparable over time and between regions since the same amount of effort is used in all predictions. For both species there is a strong seasonal pattern with peaks in September. For *O. flavescens* a secondary peak occurs in December.

Monte Carlo simulations of the number of annual strandings use the estimates displayed in Figure 3 as parameters in negative binomial probability distributions together with the scale parameter ϕ.
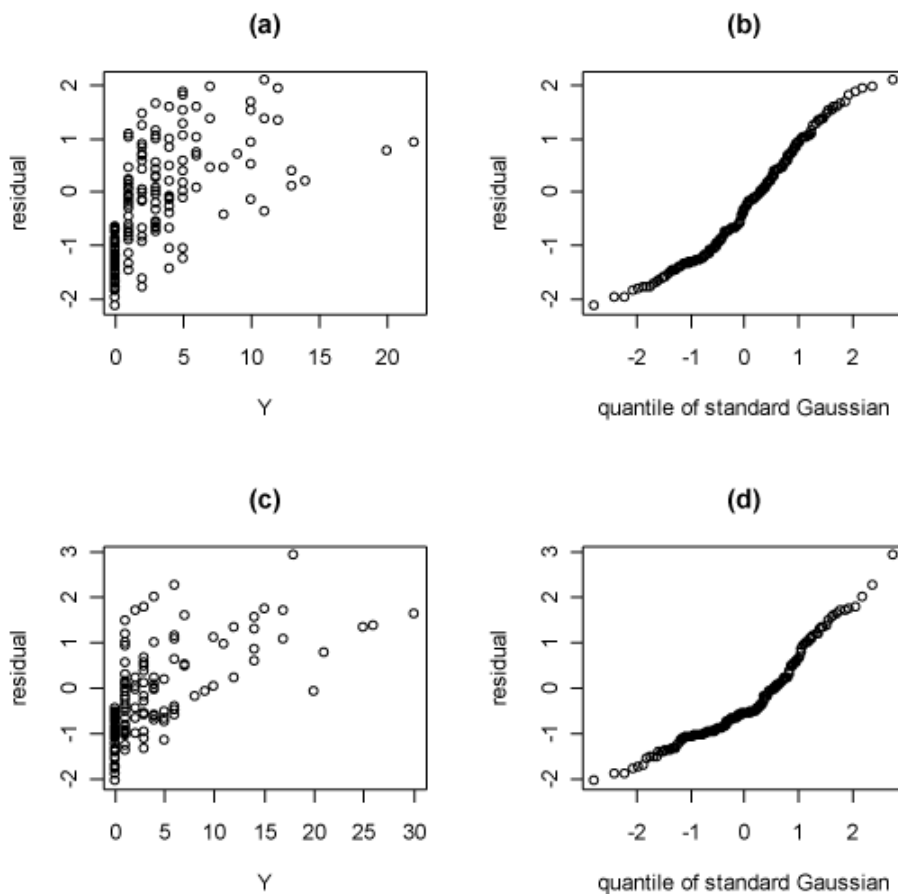
Figure 2. Residual plots for selected Model 2 with negative binomial distribution. *O. flavescens* (a, b); *A. australis* (c, d).

Results (Figure 4 and Table 4) indicate that the most common number of stranded animals is higher (aprox. 115) for *O. flavescens* than for *A. australis* (aprox. 95) but almost twice as variable for the latter (CV= 32.5%) when compared to the former (CV = 18.3%).

Discussion

The results show that GLMs are useful to formally analyze the effects of time, space and search effort on the number of stranded sea lions and seals. Since strandings are counts, the possibility to use discrete probability models (Poisson and negative binomial) is very attractive as a means of easily dealing with zeroes, and also of avoiding the need to look for appropriate transformations of the response variable *Y*.

Formal likelihood-ratio tests indicate some significant year effect for both sea lions (LRT = 20.2; p = 0.0166) and seals (LRT = 43.73; p = 1.58e-06). However, year effects present only erratic variations but no trend. In comparison, monthly variations display a stronger and much more important seasonal behavior. Furthermore, the relative increment in AIC is small when going from Model 4 to its simplified alternative Model 2 (1.9% for *O. flavescens* and 7.7% for *A. australis*) while BIC clearly favors Model 2. Hence, we believe that the extra

complexity derived from Model 4 is not worthwhile and chose Model 2 instead. Nonetheless, in other applications there might be interest in monitoring annual fluctuations in the expected number os strandings - which would require the inclusion of one or more terms in the model to account for them. At first sight it appears that in Model 2 there is no significant region effect with respect to the location being north or south of the Patos Lagoon mouth. However, this effect is included in the distinct angular effort coefficients. Depending on the region the increment per kilometer of covered coastline is statistically different with larger values in the south in comparison to the north. In Model 1, where only one angular coefficient is included, differences between regions become statistically significant and negative in all cases, confirming the pattern of fewer strandings to the north in comparison with the south. The model selection procedure, however, suggests that Model 2 provides a much better overall fit than does Model 1 (smaller AIC and BIC) and was therefore retained instead.

It is instructive to analyze (Figure 3) the discrepancies between the expected number of stranded animals, as provided by the selected models, with the observed averages per month and region. Firstly, the observed averages are mostly smaller than the expected values

estimated from the model. This is no surprise since the data were collected with unequal amounts of effort, often without covering the entire coastline. Secondly, after eliminating the effect of unequal effort which bias the observed averages, the models identify September as the month in which most strandings occur. Although this result is in line with the data of *A. australis* and with data for the northern region of *O. flavescens*, it does not agree with the data of the latter species for the southern region, which display its peak in November. Thirdly, it can be seen that the differences between observed averages and the standardized expectations provided by the selected models are much larger in the southern region, where the effort coefficients were identified to be quite larger, in comparison to the north.

The simulated numbers of annual strandings (Figure 4)

provide a clear idea of the range of possible values for a typical year and can be taken as probability distributions of deaths due to stranding to describe current uncertainties about this phenomenon in population simulation models. Since such models can be important to assist in population management, we believe this to be a very useful result to guide future conservation policies. The relatively small variability observed for *O. flavescens* (CV = 18.3%) in comparison to *A. australis* (CV = 32.5%) might be a consequence of the coastal routes adopted by sea lions (Vaz-Ferreira, 1981) such that their strandings are less affected by currents and wind directions. We recommend the maintenance of the systematically conducted monitoring program along the Rio Grande do Sul coastline in order to evaluate the evolution of mortality and the impact on population dynamics of these species.
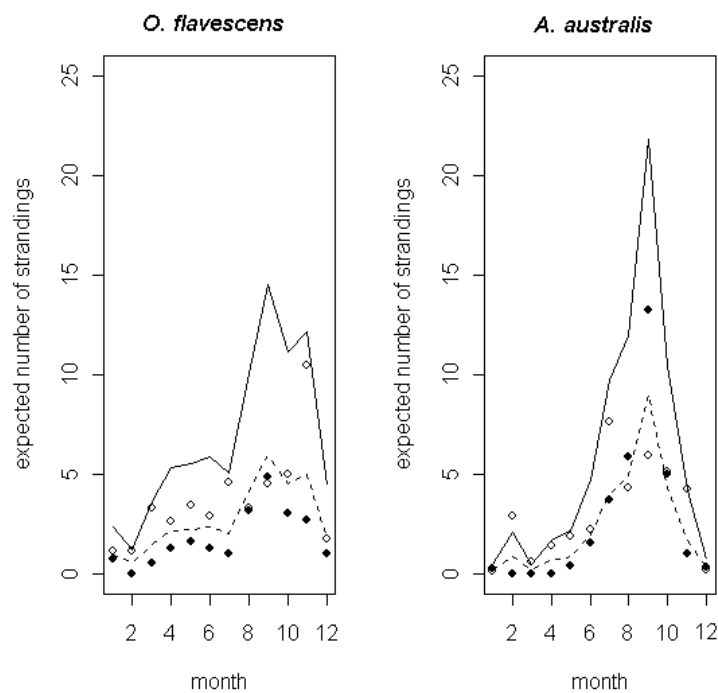

Figure 3. Expected number of strandings $\hat{\mu}$ by month for the southern region 1 (▬) and northern region 2 (▬▬), for sea lions (left) and fur seals (right) in south Brazil. The average number of observed strandings for the southern (○) and northern (●) regions are also included.

Table 4. Statistical summaries of the simulated total number of strandings for *O. flavescens* and *A. australis* south Brazil.

SUMMARY STATISTICS	O.flavescens	A. australis
Minimum	48	31
Median	113	95
Mean	114.5	99.6
Mode	~ 115	~ 95
Maximum	179	262
Stand. Dev. (CV)	21.0 (18.3%)	32.4 (32.5%)
95% C.I.	76 - 160	48 – 180

(*) Mode: center of the modal frequency class.
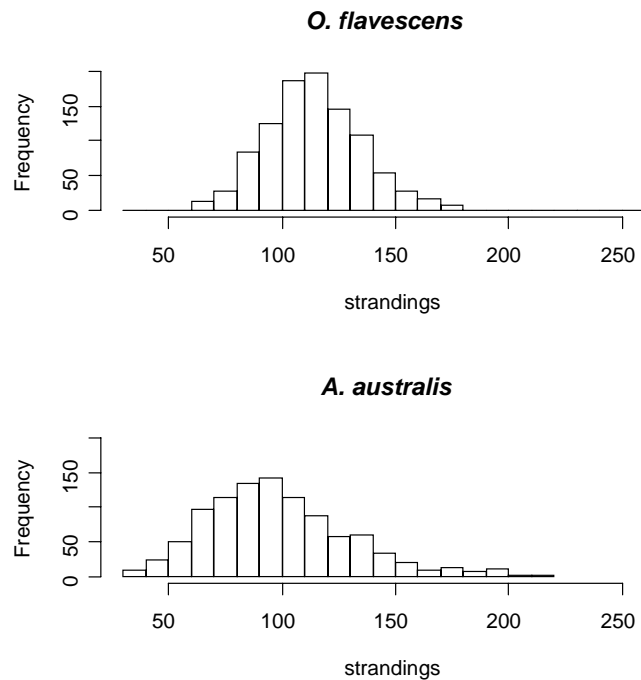
O. flavescens

A. australis

Figure 4. Distribution of 1000 Monte Carlo simulations of the numbers of annual strandings of sea lions and fur seals, south Brazil, based on model estimates from Model 2 with negative binomial distribution.

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