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POWER ANALYSIS OF POPULATION TRENDS: AN APPLICATION TO ELEPHANT SEALS OF THE FALKLANDS

Monitoring abundance is a basic requirement in conservation biology and a main source of information on population dynamics (Bart *et al.* 1998). Trends in abundance are a focus of research on Antarctic and sub-Antarctic seals (Hindell *et al.* 1994). A crucial point of trend analysis is the estimation of the likelihood of detection of a significant trend, *i.e.*, the statistical power of the analysis (Gibbs *et al.* 1998). Power calculations are important for all kinds of statistical analysis (Cohen 1988), but they seem particularly relevant when a

balance should be found among the safety of conclusions and practical, logistical, or ethical constraints (Schwagmeyer and Mock 1997).

There is an increasing appreciation of the importance of power analysis in conservation and wildlife biology (Taylor and Gerrodette 1993, Gibbs *et al.* 1998), although there is much less consensus about how to carry it out (Link and Hatfield 1991, Thomas 1997). Statistical power is defined as $1 - \beta$, with β being the probability of wrongly accepting the null hypothesis when it is actually false, *i.e.*, the probability to make a type II error. Power is, therefore, the likelihood of correctly rejecting the null hypothesis. The scope of trend analysis is to recognize sustained patterns in counts (the "signal" part of the time series), and to discriminate them from short-term fluctuations and variance due to estimation methodology (the "noise" part). The power of a monitoring program is the capability of its survey plan to detect population trends, and permit the assessment of their statistical significance (for an alternative view of the role of statistical inference see Johnson 1999). Power is influenced by many factors (Gerrodette 1987), including count error and variability, sample size, survey length, magnitude of trend to be detected (*i.e.*, in statistical terminology, the effect size), and statistical level of significance (*i.e.*, α , the probability of erroneously rejecting the null hypothesis when it is actually true, or type I error). Moreover, trends estimation is complicated by the variability of the results obtained using different methods of population estimation and fitting (James *et al.* 1996, Thomas 1996).

The analysis of trends in small populations is also complicated by the low power typical of small samples, low number of count plots, and high count variability. In such situations small trends will require long series of data to be safely detected. This is particularly unfortunate when decreasing trends are involved, because the safe detection of a small trend will be granted only after a notable reduction in the total size of the population (Forcada 2000). Hence, an accurate assessment of the status of small populations requires the estimation of the power of trend analysis.

The elephant seal (*Mirounga leonina*) population of Sea Lion Island (Falkland Islands) is a small and localized population (Galimberti and Boitani 1999). It has a particular conservation value, both locally, being the only notable breeding colony of this species in the Falklands, and globally, representing a potential conduit for gene flow between the two main populations of the South Georgia stock, South Georgia and the Valdés Peninsula (Hoelzel *et al.* 1993; unpublished data). From detailed counts carried out in 1995 and 1996, and from sparse counts carried out from 1989 to 1994, I provisionally concluded (Galimberti and Boitani 1999) that the population was almost stable. Here, I reanalyze the trend including counts for the 1997–1999 period and using robust fitting methods, and I conduct a *post-hoc* power analysis. I then run a *pre-hoc* power analysis to evaluate the trend detection capability of a long-term monitoring program based on different survey length and non-trend variability of counts.

The estimation of population size in elephant seals is complicated by the fact that, at any time, not all individuals are hauled out. Therefore, population

size is calculated using an indirect method (McCann 1985) based on the number of females hauled out during the breeding season, which may be accurately and precisely estimated from sparse daily counts by applying a mathematical model of the haul-out process (Rothery and McCann 1987, Galimberti and Sanvito 2001).

Field data were collected intensively on Sea Lion Island during the period 1995–1999. Sparse counts were available for the period 1989–1994. Details of the censusing protocol and the haul-out process modeling are available elsewhere (Galimberti and Boitani 1999, Galimberti and Sanvito 2001). The short length of the time series makes the presence of complex non-linearities unlikely. Therefore, to estimate the population trend, I fitted just two simple models, linear and exponential regression of number of females versus the year (Bart *et al.* 1998). I ran ordinary least squares (OLS) regression models with permutation tests (Manly 1991) using RT 2.0 software (Manly 1992; trial version available from <http://www.west-inc.com/west-inc.htm>). I checked the robustness of OLS regression by applying a last absolute deviation (LAD) regression, a method that is much less affected by outliers than OLS regression (Cade and Richards 1996). I ran LAD regression models using BLOSSOM software (Slauson *et al.* 1994; freely available for download from <ftp://ftp.mesc.usgs.gov/webdl>).

I ran the power analysis using the program MONITOR 6.2 (Gibbs 1995; freely available for download from <http://WWW.MP1-PWRC.USGS.GOV/powcase/monitor.html>), which uses a Monte Carlo simulation approach. The advantage of this approach is that it does not require strong assumptions about the distribution of the errors of the regression model. I checked the MONITOR results using the program TRENDS (Gerrodette 1993; freely available for download from <http://mmdshare.ucsd.edu/Trends.html>), which uses a parametric analytical approach. Results from MONITOR and TRENDS were quite similar. MONITOR produced slightly more conservative estimates of power and showed a slightly larger difference among power to detect positive and negative trends of the same magnitude. I present here the calculations from MONITOR only.

The basic constraint on the simulations was a “one plot, one count per year” survey scheme, *i.e.*, the analysis of a single, spatially localized, population surveyed once each year. I used, as mean initial value for simulation, the mean of yearly counts. Using the last count (1999) could be a valid alternative choice as starting value (Gerrodette 1993). The estimation of variability in population size due to non-trend effects is complicated because, even in absence of estimation error, the yearly estimates would not lie exactly on the trend line due to stochastic factors and short term fluctuations (Gerrodette 1987, Link and Nichols 1994). The MONITOR manual suggests the use of the variance in yearly counts as an estimate of this stochastic variability, while Thomas and Krebs (1997) showed that it should be estimated as the residual variability after trend fitting. Therefore, the MONITOR “plot variance,” and the initial coefficient of variation used by TRENDS, were calculated from the residual standard deviation of the regression model fitted to the 1989–1999 period. I

evaluated the effect of an increase in variability of counts (due to reduced accuracy of counts or increased short-term fluctuations) by increasing the standard deviation, with CV ranging from 1% to 20%. All simulations were run 10,000 times using a two-tailed 0.05 α level, and integer rounding of generated random numbers, as appropriate to count data (Gibbs 1995).

For the 1989–1996 period, the slope of the regression was close to 0 ($b = 0.32 \pm 0.78$, permutation test for slope $\neq 0$, 100,000 resamplings: $P = 0.73$). In 1997 there was an increase of 5.7%, maintained in 1998, followed by a decrease in 1999. Using all yearly values for the period 1989–1999, the slope of the regression was greater, but again not significantly different from 0 ($b = 3.25 \pm 1.73$, $P = 0.0933$). Excluding the 1997–1998 fluctuation, the increase in number of females was less than one female per year ($b = 0.96$). LAD regression confirmed this result ($b = 1.33$, $P = 0.78$). The fitting of an exponential trend gave similar results, with a non-significant 0.60% annual increase.

The residual standard deviation for the 1989–1999 counts after fitting of the linear trend was 15.50 (CV = 2.90%). I ran a Monte Carlo *post-hoc* power analysis using this value as spread of counts. Power was above 0.90 for trends of $\pm 2\%$ or more ($1 - \beta > 0.93$ for decrease trends and 0.97 for increase trends). On the contrary, it was very low for $\pm 1\%$ trends ($1 - \beta = 0.44$ for decrease trends and 0.49 for increase trends).

I ran two series of simulations to estimate *a priori* power for a long-term monitoring program. In the first series, I used the previously estimated 2.90% CV as the measure of variability of counts, and I simulated the effect of the length of the program by increasing the number of yearly surveys from 3 to 20. I fixed the target power at 0.90, which is a quite strict criterion (Cohen 1988), but seems a good balance between effectiveness and practical constraints for a small population with a potentially high extinction risk. Simulation results for a linear and an exponential trend were almost equal, and I present results of the former only. The number of years required to achieve the target power depended on the effect size, *i.e.*, the rate of the trend, in particular for small trends (Table 1). A 1% increase required 11 yr to be safely detected, and a 1% decrease 12 yr. A 5-yr-long survey safely detected only larger trends, with 6% or more increase or decrease. The detection of a negative trend required, in most cases, one yearly count more than an equivalent positive trend.

I then evaluated the effect of an increased variability of counts by systematically changing the non-trend variability in counts (CV ranging from 1% to 20%, at 1% steps). From these simulations I calculated the minimum number of yearly counts required to achieve a 0.90 power. The increase in variability of counts greatly increased the number of years required to achieve the requested power (Table 2; full power tables for this simulation are available on-line from the ESRG web site: <http://web.tiscalinet.it/esrg>). Negative 1% trends were not detectable in the 20-yr survey span when the CV of counts was higher than 6% (8% for 1% positive trends). Negative 2% trends were not detectable when the CV of counts was higher than 12% (18% for 2%

Table 1. Power of trend analysis with different size of the trend ($\pm 1\%$ – 10%) and different survey length (3–20 yr). Power values were calculated using MONITOR software with the following parameters: linear trend, 2.9% coefficient of variation due to non-trend factors, $\alpha = 0.05$, 10,000 resamplings.

	Survey length (yr)																	
	3	4	5	6	7	8	9	10	11	12	13	14	15	16	17	18	19	20
Population trend (% increase or decrease)																		
10	0.32	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
9	0.29	0.96	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
8	0.26	0.91	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
7	0.22	0.84	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
6	0.20	0.73	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
5	0.16	0.59	0.97	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
4	0.13	0.43	0.87	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
3	0.10	0.29	0.65	0.93	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
2	0.07	0.15	0.35	0.63	0.87	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
1	0.06	0.07	0.13	0.20	0.33	0.49	0.68	0.82	0.92	0.98	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00
-1	0.06	0.08	0.12	0.20	0.30	0.45	0.59	0.75	0.87	0.94	0.98	0.99	0.99	1.00	1.00	1.00	1.00	1.00
-2	0.07	0.15	0.30	0.55	0.79	0.93	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
-3	0.10	0.25	0.54	0.84	0.97	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
-4	0.12	0.37	0.76	0.97	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
-5	0.15	0.50	0.89	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
-6	0.17	0.60	0.96	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
-7	0.19	0.70	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
-8	0.23	0.78	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
-9	0.24	0.85	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
-10	0.27	0.89	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00

NOTES

Table 2. Survey length (number of years) required to achieve a 0.90 power with different size of the population trend ($\pm 1\%$ – 10%) and different non-trend variability (coefficient of variation = 1% – 20%). Power values were calculated using MONITOR software with the following parameters: linear trend, $\alpha = 0.05$, 10,000 resamplings.

	Non-trend variability (% CV)																			
	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16	17	18	19	20
Population trend (% increase or decrease)																				
10	4	4	4	5	5	5	5	6	6	6	6	6	7	7	7	7	7	8	8	8
9	4	4	4	5	5	5	6	6	6	6	7	7	7	7	7	8	8	8	8	8
8	4	4	5	5	5	5	6	6	6	7	7	7	7	8	8	8	8	9	9	9
7	4	4	5	5	5	6	6	7	7	7	7	8	8	8	9	9	9	9	10	10
6	4	4	5	5	6	6	7	7	7	8	8	8	9	9	9	10	10	10	10	11
5	4	5	5	6	6	7	7	8	8	9	9	9	10	10	10	11	11	11	12	12
4	4	5	6	6	7	7	8	9	9	10	10	11	11	11	12	12	13	13	13	14
3	4	5	6	7	8	9	10	10	10	11	12	13	13	14	14	15	15	16	16	17
2	5	7	8	9	10	11	12	13	14	15	16	16	17	18	18	19	20	20		
1	7	9	11	13	15	16	18	20												
-1	7	10	12	14	17	18														
-2	5	7	8	10	11	12	14	15	16	18	19	20								
-3	5	6	7	8	9	10	11	12	13	14	15	16	17	18	19	19	20			
-4	4	5	6	7	8	8	10	10	10	12	13	14	14	15	16	17	17	18	19	20
-5	4	5	6	6	7	8	9	9	10	11	11	12	13	13	14	15	16	16	17	18
-6	4	5	5	6	7	7	8	8	9	10	10	11	12	12	13	13	14	15	15	16
-7	4	4	5	6	6	7	7	8	8	9	10	10	11	11	12	12	13	14	14	15
-8	4	4	5	5	6	6	7	7	8	9	9	9	10	11	11	12	12	13	13	14
-9	4	4	5	5	6	6	7	7	8	8	8	9	10	10	11	11	12	12	13	13
-10	4	4	5	5	5	6	6	7	7	8	8	9	9	10	10	11	11	12	12	13

positive trends). Negative 3% trends were not detectable when CV of counts was higher than 17%.

The reanalysis of the Sea Lion Island data confirmed that the population is almost stable, but clearly showed that the current time series is not long enough to safely detect small trends in the $\pm 1\%$ range. There is some indication that the population is, in fact, increasing at less than a 1% rate, but the current data set is too small to permit safe statistical assessment of the trend. The importance of small trends should not be underestimated, even decreasing trends in the 1%–2% range may produce notable reductions of population size in the long term. The notable reduction in size of the Macquarie Island population (57.5% decline in the period 1949–1990) was due to a mere 2.1% mean annual decrease (Hindell *et al.* 1994). The population of Marion Island, which is similar in size to the Sea Lion Island population, suffered an average decline of 4.8% in the 1974–1989 period, with a slow-down to 1.9% in the last years of the period (Bester and Wilkinson 1994). This apparently small decrease rate produced a notable reduction in population size and productivity, that may easily lead to extinction in a population with an almost zero prospect of immigration from other populations. Therefore, a proper survey plan should account for this need to detect trends as small as 2%, in particular when external evidences suggest the likelihood of a decrease.

The coefficient of variation of the Sea Lion Island counts was very small if compared to values observed in surveys of large mammal populations (mean CV = 19.9%; database available at <http://www.mp1-pwrc.usgs.gov/powcase/powcase.html>). This was caused, in part, by the small fluctuations observed in the population, but also by the very accurate counts repeated during the whole length of the breeding season. By modeling the haul-out process (Galimberti and Sanvito 2001), we demonstrated that a very good estimate of the total number of breeding females could be obtained with a small number of daily counts. If counts are carried out close to the day of peak presence on land of females (*i.e.*, in a 1-wk period centered on the peak) even a single count guarantees an estimation error for the total number of breeding females within $\pm 2\%$. Therefore, even with a relaxed censusing protocol, it should be possible to keep the total variation (count error plus fluctuations) within a 5% limit. This means that a $\pm 2\%$ trend could be detected in 10 yr, and that even a $\pm 1\%$ trend could be detected in 20 yr. The planning of a survey for southern elephant seals populations should, therefore, be based on a two-stage process. Firstly, for at least one breeding season, frequent counts (weakly at least) should be carried out to have enough data point to obtain a good estimate of the peak haul-out day, and a good model of the haul-out process. This model should possibly be cross-validated with data from the following season. Then a yearly survey may be carried out with a modest effort, because just one or a few counts carried out close to the peak will be enough to obtain a good estimate of the total number of females.

In conclusion, power analysis seems to be a useful tool for the planning of monitoring programs for seal populations. On the other side, the intrinsic and extrinsic limitations of power analysis need to be taken into account. Power

analysis, both with Monte Carlo and analytical methods, produces approximate results (Gerrodette 1991). Exact procedures for power estimation are available for just a very small number of methods and in specialized software only (Metha and Pathel 1998). An advantage of a randomization approach is that the precision of the estimate may be increased as much as desired by simply increasing the number of random resamples (Manly 1991). The results of power analysis are affected by many factors, including the choice of the trend model (linear, exponential, or customized to account for specific non-linearities). Moreover, the trend estimation can be affected by autocorrelation among yearly values, that produce biased regression standard errors and power estimates. Unfortunately, the estimation of autocorrelation, and of the time lag at which autocorrelation is maximum, is limited by the length of the time series (Legendre and Legendre 1998). The maximum allowable lag is about $n/4$ (where n is the number of yearly values) and, therefore, with short series only the first order autocorrelation can be calculated. The econometric literature offers various methods to improve the estimate of trend error in case of autocorrelation (*e.g.*, Newey and West 1987), which are implemented in both commercial (*e.g.*, STATA, <http://www.stata.com>) and free software (*e.g.*, LIM-DEP, <http://www.limdep.com>). Notwithstanding these limitations, power analysis permits a quantitative assessment of the likelihood of trend detection, and a valuable comparison of the pros and cons of alternative survey designs. The wide availability of software to carry out power analysis (Thomas and Krebs 1997) should put the calculation of power of trend detection in the toolkit of any researcher interested in abundance estimation.

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