Indices of Condition for Small Mammals

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Abstract

Estimates of body condition in mammals may be constructed from measures of skeletal size and body mass. We illustrate the methodology for doing this using data from two populations of feral house mice (*Mus domesticus*) in Australia, and point out an erroneous method that has commonly been used in the literature. Indices of condition for individual house mice were not correlated with the fat content of their carcasses. Indices of condition for house mice have a relatively low repeatability because of variation from day to day in body mass and because of variation in length measurements taken by different observers. Bias in measurements among observers must be eliminated to make indices of condition from live animals useful.

Introduction

Population biologists are often interested in measuring the body condition of small mammals non-destructively. The normal method for doing this is to measure skeletal size and body mass (LeCren 1951; Bailey 1968; Angerbjörn 1986). Skeletal size may be measured as total length, body length, hind-foot length, or any measure of the skeleton that can be taken on live animals. In this paper we discuss how to use data of this type to construct an index of condition.

A common error in estimating condition is to apply the general principles of scaling theory to these body measurements. Theory suggests from simple geometry that body mass should be related to the cube of length, if mammals are like simple cylinders. By dividing body mass by a measure of length cubed, we can thus get a simple index of condition. We illustrate here with data on feral house mice (*Mus domesticus*) in Australia why this is not an appropriate measure of condition.

Methods

House mice were live-trapped at two farms in the Victorian mallee at Walpeup during 1983–85 (Singleton 1989) and at 7 farms on the Darling Downs, Queensland, during 1992 and 1993. Each individual was weighed with a Pesola spring balance to the nearest 0·1 g and head–body length was measured in the field to the nearest millimetre. Similar methods of measurement for different observers were achieved by periodic repetitive measurements of the same individual mice by different observers. Repeatability estimates for mass, length and condition were calculated as in Krebs (1989, p. 453). Live-trapping procedures for house mice utilised standard grid trapping with Longworth live-traps. These procedures are described in detail in Singleton (1989).

Regression analysis was carried out with SYSTAT (Wilkinson 1988). We used multivariate general linear hypothesis routines to test for equality of slopes and to compute regression coefficients. Standard regressions were used in preference to geometric mean regressions (GMR) because there were no serious measurement errors in lengths or mass and we wished to use the regressions to predict mass from a given length (Krebs 1989, p. 463).

From May 1983 to November 1985, mice were collected at approximately 6-weekly intervals from the Victorian mallee. The stomachs of these mice were removed for analysis of their diet (see Tann *et al.* 1991 for details) and the carcasses used for fat analysis. Each carcass was frozen for transport to the laboratory. After thawing, the carcass was dried at 60°C for 14–16 h in a vacuum oven at 25 inches of Hg, cooled in a desiccator, weighed and then pulverised in a blender for 15 s. The contents were then transferred to a pre-weighed cellulose, single-thickness, 22×80-mm extraction thimble (Whatman, U.K.). Thimbles were inserted in a soxhlet apparatus and fat was extracted for 5 h with technical-grade carbon tetrachloride. After the extraction, thimbles were dried at 60°C for 14–16 h, cooled in a desiccator and weighed. The body fat was measured as a percentage of dry body mass. Pregnant females were excluded from fat analyses.

Results

The estimation of condition involves three steps: (1) estimating the regression between skeletal size (X) and body mass (Y) for the population; (2) using this regression to predict body mass from observed skeletal size for each individual; and (3) estimating the condition of each individual from the ratio of observed body mass to predicted body mass.

The first step of estimating a regression for the propulation under study should be done on a large data set including several years and areas if possible and including both sexes and all age classes. For house mice we have calculated separate regressions for the two study areas. There is a slight but significant curvilinear relationship between mass and body length in house mice (Fig. 1*a*). A second-degree polynomial offered a significantly better fit than a straight line for these regressions ($P < 0.001$). A log–log regression straightens the regression and stabilises the variance about the regression (Fig. 1*b*). We have therefore used log–log regressions in our analysis.

We pooled all our data at each site. For the Victorian mallee we obtained the regression

$$\log_e M = 2 \cdot 5689 \ L - 8 \cdot 5650 \qquad (r = 0 \cdot 907, \ n = 5888)$$

where L is body length. For the Darling Downs we obtained the regression

$$\log_e M = 2 \cdot 4300 \ L - 8 \cdot 0008 \qquad (r = 0 \cdot 921, \ n = 2407).$$
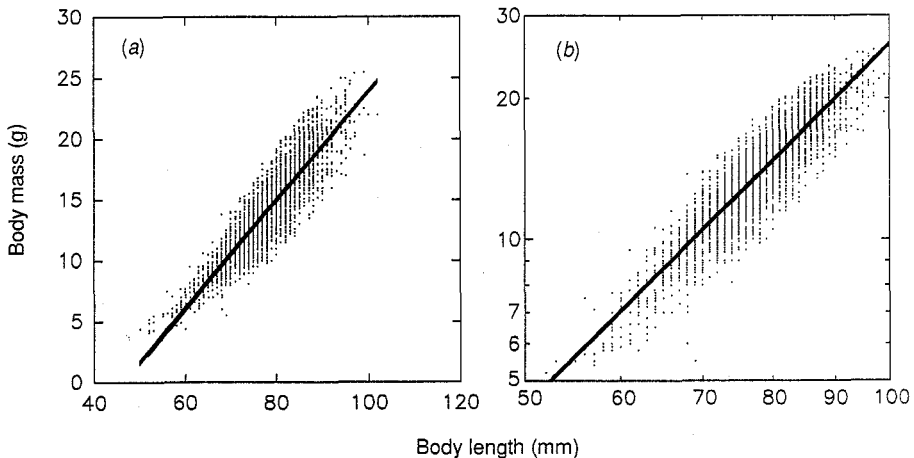

Fig. 1. Relationship between head and body length (nose to anus) and body mass in feral house mice from the Darling Downs, Queensland in 1992–93. A total of 2407 mice were measured. There was no significant difference in slope between the regression for males and that for females ($P = 0 \cdot 97$). All pregnant females were omitted. (*a*) Linear regression: $Y = 0 \cdot 4019X - 17 \cdot 88$, $r = 0 \cdot 915$, $n = 2407$. (*b*) Curvilinear regression ($\log_e$–$\log_e$: $Y = 2 \cdot 4300X - 8 \cdot 0008$, $r = 0 \cdot 921$. Note that the linear regression overestimates body mass for medium-sized mice and underestimates it for large-sized mice.

These regressions differ significantly from each other in slope. Both these regressions include both sexes (3532 males and 2356 females for Victoria; 1511 males and 896 females for Darling Downs). Using analysis of covariance, we tested the assumption of equality of slopes (Wilkinson 1988, p. 528) for the two sexes (excluding pregnant females) and found that the slopes did not differ between sexes for either study area ($P = 0.51$ for Victoria, $P = 0.97$ for Darling Downs). The elevation of the regression line for each sex differed significantly in both areas ($P < 0.05$). Female house mice were very slightly lighter than male house mice (on average 13·12 v. 13·31 g, adjusted means).

The decision of what to include in a regression should be based on the groups in the population you wish to compare. If you wish to compare males and females, you should use a common regression. If you do not wish to compare the two sexes, and they differ in their regression lines, you can compute a different regression for each sex. For house mice we wished to compare males and females and therefore we computed a common regression for all individuals.

Given the regression line, we can estimate body mass and condition. An example will illustrate this procedure. A male house mouse from the Victorian mallee had body length 94 mm and weighed 19·9 g. From the above regression for this region:

$$\text{Predicted body mass} = \exp \, [2 \cdot 5689 \, (\log_e 94) - 8 \cdot 5650]$$
$$= 22 \cdot 3 \text{ g}$$

For this individual,

$$\text{Index of condition} = \text{observed mass/predicted mass}$$
$$= 19 \cdot 9 / 22 \cdot 3$$
$$= 0 \cdot 890$$

The average individual should have an index of condition of 1·0, and this particular mouse is about 11% below average in condition. For any particular population, these indices may be averaged and analysed statistically in the usual ways. The standard deviation of the condition index of house mice is 0·0391 and the distribution is nearly normal in shape. A power curve for the sample sizes needed to pick up variations in the index of condition for house mice is shown in Fig. 2. If you wish to detect with high probability a difference of 0·05 in mean condition between two populations, you will require a sample size of more than 15 individuals from each population.
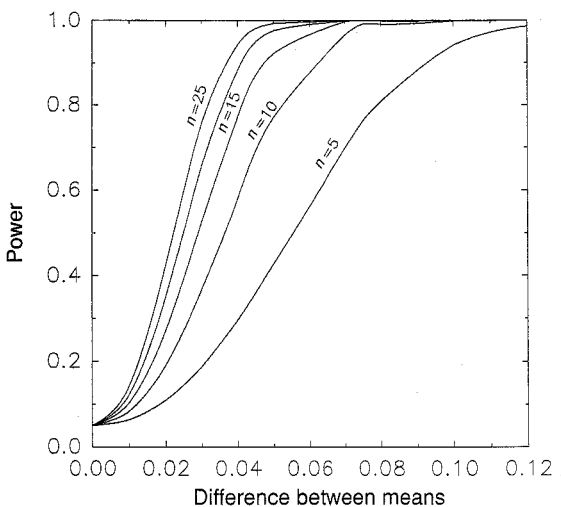

Fig. 2. Statistical power curves for the difference between mean indices of condition for two groups of house mice for sample sizes from $n = 5$ to $n = 25$. All assume $\alpha = 0 \cdot 05$, s.d. $= 0 \cdot 0391$ (data from Victorian mallee) for the index of condition. Note that sample size is the sample size needed for each of the two groups being compared.

The scaling theory approach to estimating condition assumes a regression of the form:

$$Y = a\ L^3$$

or
$$\log Y = \log a + 3 \cdot 0 \log L$$

so that a prediction of this model is that the slope of the log–log regression of mass on length is $3 \cdot 0$. For house mice this is not correct. The slope of this regression is $2 \cdot 569 \pm 0 \cdot 016$ (s.e.) for Victoria and $2 \cdot 449 \pm 0 \cdot 021$ for the Darling Downs. Given this slope, an index of condition calculated from the scaling model with exponent 3 would progressively underestimate condition as animals become larger. To illustrate this, consider two mice from the Victorian mallee of average size so that their indices of condition are $1 \cdot 0$:

Mouse	Body length	Body mass	Index of condition
A	75 mm	12·51 g	1·00
B	90 mm	19·98 g	1·00

These individuals are on the regression line given above, so they represent mice that differ in overall size but have identical body condition. If we calculate their index of condition from scaling theory we obtain an incorrect conclusion that the larger mouse is in relatively poor condition:

| Mouse	Scaling index	Relative condition
A	$2 \cdot 965 \times 10^{-5}$	$+3 \cdot 9\%$
B	$2 \cdot 741 \times 10^{-5}$	$-3 \cdot 9\%$

If we use the geometric mean regression (Krebs 1989, p. 463) in place of the standard regression to describe these relationships, the slopes are closer to the theoretical $3 \cdot 0$ predicted by scaling theory but remain significantly lower. For the Darling Downs data ($n = 2407$) the slope of the GMR is $2 \cdot 685 \pm 0 \cdot 021$ (s.e.) and for the Victorian data ($n = 5888$) the slope is $2 \cdot 929 \pm 0 \cdot 014$ (s.e.).

Indices of condition are expected to provide a general measure of the health of individual animals. Variation in body mass due to stomach contents or time spent in a live-trap will add variance to this measure. One way to test the critical assumption of condition indices is to measure these for individuals that are subsequently analysed for fat content. The
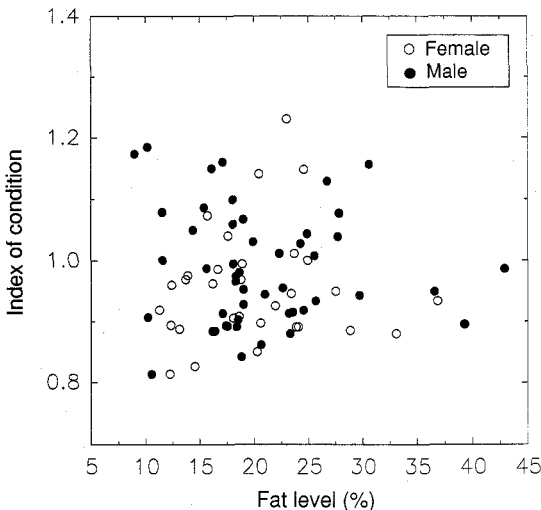
Fig. 3. Relationship between total body fat content (%) and the index of condition for 49 male and 31 female house mice collected in the Victorian mallee. Pregnant females were excluded. Spearman's $r = 0 \cdot 08$ for females, $-0 \cdot 07$ for males, both non-significant statistically.

assumption is that fat content is a good measure of general condition, and thus there should be a high correlation between the index of condition as defined above and the fat content of the whole carcass. Fat measurements were available for 80 house mice from which data on body mass and body length were also taken. The relationship between the fat content of these individuals and their index of condition is shown in Fig. 3. There is no correlation at all between these variables, contrary to expectation. To determine whether this lack of correlation was due to individual variation in fat and body measurements, we used a larger data set from the Victorian mallee and plotted, for the entire sample of 602 mice, the mean percentage fat for each week sampled against the mean index of condition. It can be seen from Fig. 4 that there is also no relationship between these mean values, corroborating the results from the smaller sample of individuals used in Fig. 3.
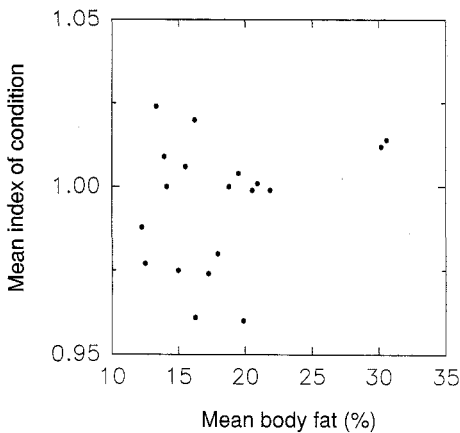
Fig. 4. Relationship between mean body fat content (%) and the mean index of condition for 19 samples from 1983 to 1985 in the Victorian mallee. Each point represents one week's sample. Total sample size for fat determination was 602 mice and total sample size for the index of condition was 2973 mice. Spearman's $r = 0.00$, $P > 0.50$.

Discussion

There is considerable confusion in the literature on mammals about the proper method of calculating an index of condition from skeletal and mass data, stretching back to Quetelet (1869). LeCren (1951) discussed this problem for fish in which the rationale for an index of condition based on scaling (the cube law) is more nearly correct. He calls the type of index we have used a *relative condition factor*. Bailey (1968) found for his skeletal measures of cottontail rabbits (*Sylvilagus floridanus*) that the slope of the regression of log(length) on weight was exactly 3. Others seem to have adopted his slope value without testing for it directly. It is unlikely that for most measures of skeletal size the slope will be exactly 3. Angerbjörn (1986) used a square of the hindfoot measurement for *Lepus timidus* to calculate an index of condition without validating the exponent 2. Bakker and Main (1980) used a complex method of measuring condition in the quokka (*Setonix brachyurus*) by taking the cube root of body mass and relating it to skeletal measurements. This method again presupposes the validity of the cube law of scaling, and has the potential for the type of errors discussed above in which animals of different size appear to differ in condition when in fact they are equal. Garrow (1983) discusses this problem for humans and points out the same potential errors.

There is a basic assumption underlying all measures of condition that in fact they measure some attribute of an individual that can be labelled 'condition'. The presumption in an ecological context is that individuals of higher condition will have higher reproductive rates, lower mortality rates, and be able to cope with environmental stresses, starvation, or cold temperature more easily (Brochu *et al.* 1988). We can find no literature that assesses this critical assumption. Our finding that the index of condition is not correlated at all with

fat content in house mice in disturbing. It can mean that fat levels are not a good measure of body condition in these small rodents that have high metabolic turnover. Alternatively, it could mean that variations in observed body mass are determined by variables like gut contents too often to be useful in field studies. It is important to determine which of these alternatives is correct before we can confidently recommend the use of condition indices based on skeletal measurements.

One way to test the variable body weight alternative is to determine the index of condition for individuals at frequent intervals over a short period and to calculate repeatability of the index on the assumption that it should not change from day to day. We obtained independent mass and length measurements for three groups of house mice from the Darling Downs and the repeatability of the measurements is given in Table 1. The repeatability of the index of condition ranged from 0·15 to 0·78 and was most highly dependent on the repeatability of body mass. We conclude that there is relatively low repeatability of the index of condition in house mice and this results from large variation in body mass from day to day (possibly due to time spent in the live-trap) and from variation in the measurement of body length.

Errors of measurement are minimal for body mass in our data, but there could be significant inter-observer variation in the measurement of skeletal size. In any study that uses indices of condition it is important to standardise the taking of these measurements. Errors of 1 mm in measuring body length in house mice result in approximately a 1% change in the estimated index of condition. If errors of measurement are unbiased only the variance of the index will be affected, but if there is a systematic bias (one observer consistently measuring longer lengths than another) the value of any index will be compromised.

Table 1. **Repeatability of house mouse measurements of body mass, length and index of condition for three samples from the Darling Downs, 1992–93**

95% confidence limits in parentheses

| Date | Sample size | | Repeatability | |
		Body mass	Body length	Index of condition
November 1992	39	0·955 (0·911–0·977)	0·951 (0·919–0·979)	0·389 (0·095–0·622)
April 1993	33	0·936 (0·876–0·968)	0·978 (0·956–0·989)	0·152 (0·000–0·465)
May 1993	27	0·985 (0·968–0·993)	0·963 (0·922–0·983)	0·776 (0·570–0·891)

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