Sex-bias coalitions results

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8 2 2022

To estimate the overall predicted probabilities of coalitions by females only, males only, or both sexes, we used multinomial models. Contrary to our first prediction, female only coalitions were not more likely than male only coalitions (Fig. 1, Model 1). In fact, the probability of female only coalitions (median = 0.24, 90% credible interval = 0.02 - 0.58) was virtually the same as the probability of male only coalitions (0.31, 0.02 - 0.64) and the most likely state was coalitions by both sexes (0.36, 0.11 - 0.63). Thus, only 44% of the posterior probability supported our prediction of female only coalitions being more likely than male only coalitions, and none of the three possible states was greater than chance (% of posterior > 0.33 for females only: 35%; males only: 45%; both sexes: 58%). The phylogenetic signal was weak (median λ = 0.41, 90% CI = 0.22 - 0.61). In other words, conditional on having any coalitions at all, the typical extant mammal is just as likely to have female only, male only, or both sex coalitions.

This general pattern did not change appreciably when comparing species living in mixed-sex groups (n = 53 species) with sex-segregated ones (n = 4 species; Model 2), or primates (n= 38 species) with non-primates (n= 19 species; Model 3). Specifically, the odds of female only coalitions and male only coalitions were essentially the same in sex-segregated species compared to mixed-sex species (females only: median odds ratio = 1.05, 90% CI = 0.32 - 2.02, probability OR>1 = 54%; males only: median OR = 1.03, 90% CI = 0.31 - 2, probability OR>1 = 52%) or primates compared to non-primates (females only: median odds ratio = 0.99, 90% CI = 0.33 - 1.97, probability OR>1 = 49%; males only: median OR = 1, 90% CI = 0.31 - 1.95, probability OR>1 = 50%). We therefore did not stratify our subsequent analyses by these variables.

To test socio-ecological predictions about female coalitions, we combined the three categories “females only”, “males only” and “both sexes” into a binary variable for the presence (“females only”, “both sexes”) or absence (“males only”) of female coalitions; philopatry was also re-coded as a binary variable indicating presence (females or both sexes philopatric) or absence (males or neither sex philopatric) of female philopatry. To test predictions about male coalitions, we analogously re-coded male coalitions as present (“males only”, “both sexes”) or absent (“females only”) and male philopatry as present (males or both sexes philopatric) or absent (females or neither sex philopatric). We then ran binomial models on the presence of female coalitions (Model 4), including food defensibility (yes/no) and female philopatry (yes/no) as predictors, and one on the presence of male coalitions (Model 5), including sexual dimorphism and male philopatry (yes/no) as predictors. As a robustness check, we also modeled each of these competing causes on its own.

The probability of female coalitions was not more likely in species with defensible food resources compared to species with non-defensible food resources (OR = 0.93, 90% CI = 0.33 - 1.87, probability OR>1 = 45%), or in species with female philopatry compared to female dispersal (OR = 1.03, 90% CI = 0.31 - 2, probability OR>1 = 52%). Likewise, the probability of male coalitions was only marginally higher in sexually dimorphic species (OR for 1SD change in dimorphism = 0.89, 90% CI = 0.29 - 1.64, probability OR>1 = 40%), and virtually the same whether males were philopatric or dispersed (OR = 1.09, 90% CI = 0.29 - 2.1, probability OR>1 = 56%). These inferences did not change when considering each predictor in a model on its own (see Supplement). Thus, the probability of female or male coalitions was not strongly associated with our predictors.