Title:

Formal Modelling of Predator Preferences using Molecular Gut-Content Analysis

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Summary

1. The literature on modelling the prey selection a predator includes many intuitive indices, few of which have both reasonable statistical justification and tractable asymptotic properties.

2. Here, we provide a simple model that meets both of these criteria, while extending previous work to include an array of multiple species and time points.

3. Further, we apply the Expectation-Maximisation algorithm to cases where exact counts of the number of prey species eaten in a particular timer period is not observed.

4. We conduct a simulation study to demonstrate the accuracy of our method, and a real dataset, collected on wolf spiders using molecular gut-content analysis, illustrates how to apply our methods.

1 Introduction

The indices most commonly used to estimate a predator's food preferences, or selectivity, are relatively old (Ivlev, 1964; Jacobs, 1974; Chesson, 1978; Strauss, 1979; Vanderploeg and Scavia, 1979; Chesson, 1983), and yet many applied papers continue to use them; a quick search of papers published in 2014 returns hundreds of publications that cite these fundamental papers, a few being Clements et al. (2014); Hansen and Beauchamp (2014); Hellström et al. (2014); Lyngdoh et al. (2014); Madduppa et al. (2014). These indices, though intuitive, lack the statistical rigour of a model, focus on a snapshot in time, and rarely allow more than one prey species to be considered (Lechowicz, 1982). We propose an intuitive statistical model to determine and statistically test differences in a predators' prey preferences across an array of time points and between multiple prey species.

A comprehensive overview by Lechowicz (1982), which was later summarized by Manly et al. (1992), details the benefits and faults of the most popular indices. According to these reviews, a majority of the indices give comparable results, save Strauss' linear index L, despite the fact that most of the methods differ by range and linearity of response. While Lechowicz (1982) recommends one index, E^* by Vanderploeg and Scavia (1979) as the "single best," albeit imperfect, index, Manly et al. (1992) instead take the approach of excluding the subset of indices which do not "estimate any biologically meaningful value." Lechowicz (1982) recommends the index E^*, an element of the Manly et al. (1992) suggested indices, because the index value 0 denotes random feeding, the index has a range restricted to $[-1, 1]$ (though $E^* = 1$ is nigh impossible), and because the index is based on the predator's choice of prey as a function of both the availability of the prey as well as the number of available prey types (assumed known). The downside to this index is its lack of reasonable statistical properties (Lechowicz, 1982), thus making the computation of standard errors, and hypothesis testing difficult. This is, in fact, a common fault amongst most of the indices.

To encourage more formal statistical inference, and simultaneously generalise predators' selectivity to animal resource selection, Manly et al. (1992) proposed the use of generalised linear models (GLM). The well established literature on GLMs allows for hypothesis testing

to replace the indices, by estimating the proportion of eaten prey species to that which is available while using environmental variables as predictors. The model we present here, while restricted to predators' preferences, is a compromise between these two extremes, indices and GLMs. Our model offers formal hypothesis testing and inference similar to the GLMs of Manly et al. (1992), but also provides meaningful single number summaries of the predator's dietary preferences. To do this, we estimate the rate at which a predator consumes the prey of interest instead of estimating the proportion of consumed to available prey. An outcome of our model, is that we give up the somewhat arbitrary preference for an index to have the range $[-1, 1]$, and random feeding, now denoted by 1, is formally testable across time points and across prey species.

Our model enables formal hypothesis testing and statistical inference, while being general enough to perform statistical tests across multiple species and time points. This provides researchers a more detailed analysis of the predator's feeding preferences. Further, because our model is based on underlying Poisson distributions, members of the well studied exponential family, we are able to estimate the parameters of interest even when exact tallies of the number of each prey species eaten within any given time period is not observed. Instead, we rely on the researcher being able to detect prey DNA within the predator's gut (Schmidt et al., 2014; Raso et al., 2014; Madduppa et al., 2014) and make a simple binary conclusion: this predator ate some of that prey species during this time period, or did not.

This paper is organized as follows. Section 2 describes our statistical model, for both fully observed count data, and for the non-observed count data for which we use the Expectation-Maximization (EM) algorithm, and the statistical tests used to make statements about the population parameters of interest. In section 3, we offer a simulation study that demonstrates the accuracy of our methods. Section 4 provides a real dataset, which investigates the eating preferences of wolf spiders (Araneae: Lycosidae), found in the Berea College Forest in Madison County, Kentucky, USA, to demonstrate how those interested in assessing trophic interactions with gut-content analyses could apply our methods. A brief discussion concludes the paper in section 5. Alongside our model, we offer an `R` (Core Team, 2014) package named `spiders` that fits all the methods discussed.

2 Methods

2.1 Data

We assume data are collected in the following manner. Traps are dispersed, for T time periods, throughout the habitat of the predator and prey of interest. Prey species, indexed by $s \in \{1, \ldots, S\}$, are collected in the traps and counted at each time period. The number of prey species s the predator will encounter on average during time period $t \in \{1, \ldots, T\}$ are considered random draws from a Poisson distribution with rate parameter γ_{st}. We further assume that the number of prey species found in the gut of the similarly trapped predators follows a Poisson distribution with rate λ_{st}. Here, the parameter λ_{st} represents the rate at which the predator ate prey species s during time period t. By modeling λ_{st} and γ_{st} we are able to test claims about a predator's eating preferences.

The use of Poisson distributions make the following implicit assumptions: 1) traps inde-

pendently catch the prey species of interest, 2) predators eat independent of each other.

We denote the number of predators and the number of prey species caught, in each time period t, by J_t and I_t, respectively. Let $X_{jst} \overset{\text{iid}}{\sim} \mathcal{P}(\lambda_{st})$ represent the number of prey species s that predator j ate during occurrence t, where $j \in \{1, \ldots, J_t\}$. Let $Y_{ist} \overset{\text{iid}}{\sim} \mathcal{P}(\gamma_{st})$ represent the number of prey species s found in trap i during occurrence t, $i \in \{1, \ldots, I_t\}$. Formal statistical statements about the relative magnitudes of the parameters $\boldsymbol{\lambda}$ and $\boldsymbol{\gamma}$ offer insights to the relative rates at which predators eat particular prey species.

We consider five variations on the relative magnitude of $\lambda_{st}/\gamma_{st} = c_{st}$. These five hypotheses each allow c_{st} to vary by time, prey species, both, or neither. Because the five hypotheses are nested, a natural testing order is suggested in Figure 1.

1. $c_{st} = 1$
2. $c_{st} = c$
3. $c_{st} = c_s$
4. $c_{st} = c_t$
5. $c_{st} = c_{st}$

The first hypothesis states that predators and traps sample all prey species at the same rate. One imagines this is the case if the predator simply eats prey which comes within its reach, thus suggesting no selection for a particular prey item. The second states that predators sample prey proportionally across all time periods. The third hypothesis states that predators sample different prey species at different rates, but each rate is steady across time. This implies that the predator expresses preferences for one prey species over another, but is unresponsive to changes due to time. Conversely, the forth hypothesis implies that each prey species is sampled similarly within each time period, while the rates across time are allowed to change. The fifth hypothesis assumes a predator's selection varies by both time and prey species. This would make sense if environmental and biological variables, such as weather, prey availability, and/or palatability were affecting predators' selection strategies.

[Figure 1 about here.]

2.2 Fully Observed Count Data

The likelihood function that allows for estimation of these parameters is as follows. Since we assume X_{jst} is independent of Y_{ist} we can simply multiply the respective Poisson probability density functions, and then form products over all s, t to obtain the likelihood.

$$L(x_{jst}, y_{ist}|\boldsymbol{\lambda}, \boldsymbol{\gamma}) = \prod_{t=1}^{T} \prod_{s=1}^{S} \left\{ \prod_{j=1}^{J_t} f_X(x_{jst}|\boldsymbol{\lambda}) \prod_{i=1}^{I_t} f_Y(y_{ist}|\boldsymbol{\gamma}) \right\}. \tag{1}$$

Writing all five hypotheses as $\lambda_{st} = c_{st}\gamma_{st}$, we can, in the simplest cases, find analytic solutions for the maximum likelihood estimates of c_{st} and γ_{st}. Under the hypothesis $c_{st} = 1$,

and when the data are balanced $J_t = J$, $I_t = I$, and $c_{st} = c$ analytic solutions exist. Namely, these solutions are

$$\hat{\gamma}_{st} = \frac{X_{\cdot st} + Y_{\cdot st}}{J_t + I_t}, \quad \text{and} \quad \hat{c} = \frac{I \sum_{s,t} X_{\cdot st}}{J \sum_{s,t} Y_{\cdot st}}, \quad \hat{\gamma}_{st} = \frac{X_{\cdot st} + Y_{\cdot st}}{I\left(\frac{\sum_{st} X_{\cdot st}}{\sum_{st} Y_{\cdot st}} + 1\right)}$$

respectively, where $X_{\cdot st} = \sum_{j=1}^{J_t} X_{jst}$ and $Y_{\cdot st} = \sum_{i=1}^{I_t} Y_{ist}$.

In all other cases, analytic solutions are not readily available and instead we rely on the fact that the log-likelihood $l(\boldsymbol{\lambda}, \boldsymbol{\gamma}) = \log L$ is concave. We maximize the log-likelihood, using coordinate descent (Luo and Tseng, 1992), by iteratively solving partial derivatives of l, with respect to c_{st} and γ_{st}, set equal to zero

$$\hat{c} = \frac{\sum_{s,t} X_{\cdot st}}{\sum_t J_t \sum_s \gamma_{st}}, \quad \hat{c}_t = \frac{\sum_s X_{\cdot st}}{J_t \sum_s \gamma_{st}}, \text{ or } \quad \hat{c}_s = \frac{\sum_t X_{\cdot st}}{\sum_t J_t \gamma_{st}}, \text{ and } \quad \hat{\gamma}_{st} = \frac{X_{\cdot st} + Y_{\cdot st}}{J_t c_{st} + I_t}.$$

2.3 Unobserved Counts

In many applications, such as DNA-based gut-content anslysis, it is not possible to count the number of individuals of each prey species that are in a predator's gut. Instead, it is only possible to detect whether or not a predator consumed the prey species during a given time period, based on the rate at which prey DNA decays in the predator gut (Greenstone et al., 2013). In this case we can still make inference about the predators' preferences for the different prey species by using the Expectation-Maximization (EM) algorithm to compute maximum likelihood estimates.

We denote the binary random variable indicating that the j^{th} predator did in fact eat at least one individual of prey species s in time period t by $Z_{jst} = 1(X_{jst} > 0)$. The variables are independent Bernoulli observations with success probability $p_{st} = P(Z_{jst} = 1) = 1 - \exp\{-\lambda_{st}\}$. Despite not observing X_{jst}, we can compute maximum likelihood estimates of the parameters $\boldsymbol{\lambda}, \boldsymbol{\gamma}$ through the EM algorithm using the complete data log-likelihood

$$l_{comp}(\boldsymbol{\lambda}, \boldsymbol{\gamma}) = \log f_{X,Y,Z}(\boldsymbol{x}, \boldsymbol{y}, \boldsymbol{z}|\boldsymbol{\lambda}, \boldsymbol{\gamma}) = \sum_{s=1}^{S} \sum_{t=1}^{T} \left[\sum_{j=1}^{J_t} \log f_{X,Z}(x_{jst}, z_{jst}|\boldsymbol{\lambda}) + \sum_{i=1}^{I_t} \log f_Y(y_{jst}|\boldsymbol{\gamma}) \right].$$

The density of Y_{jst} is exactly as in section 2.2 and so we focus on deriving the joint density of X_{jst} and Z_{jst}. With the distribution of Z_{jst} given above, we can compute $f_{X,Z}(x_{jst}, z_{jst}|\boldsymbol{\lambda})$ by noting that $X_{jst} = 0$ with probability 1 if $Z_{jst} = 0$, and that $[X_{jst}|Z_{jst} = 0]$ has a truncated Poisson distribution with density

$$f_{X|Y,Z,\boldsymbol{\lambda},\boldsymbol{\gamma}}(x_{jst}|z_{jst}) = \frac{\exp\{-\lambda_{st}\}\lambda_{st}^{x_{jst}}}{(1 - \exp\{-\lambda_{st}\})x_{jst}!} 1(x_{jst} > 0)$$

and expected value

$$\mathbb{E}_{X|Y,Z} X_{jst} = \frac{\lambda_{st} \exp\{\lambda_{st}\}}{\exp\{\lambda_{st}\} - 1}.$$

The joint density of X_{jst}, Z_{jst} is then

$$f_{X,Z|\boldsymbol{\lambda}}(x_{jst}, z_{jst}) = \begin{cases} \exp\{-\lambda_{st}\}, & x_{jst} = 0 \text{ and } z_{jst} = 0 \\ \frac{\exp\{-\lambda_{st}\}\lambda_{st}^{x_{jst}}}{x_{jst}!}, & x_{jst} > 0 \text{ and } z_{jst} = 1 \\ 0 & \text{otherwise} \end{cases}.$$

The EM algorithm works by iterating two steps, the E-step and M-step, until the optimum is reached (Dempster et al., 1977; McLachlan and Krishnan, 2007). Let k index the iterations in the EM algorithm so that $\boldsymbol{\lambda}^{(k)}$ and $\boldsymbol{\gamma}^{(k)}$ denote the estimates computed on the k^{th} M-step. The E-step consists of computing the expectation of l_{comp} with respect to the conditional distribution of X given the current estimates of the parameters

$$Q^{(k)}(\boldsymbol{\lambda}, \boldsymbol{\gamma}) = \mathbb{E}_{X|Y,Z,\boldsymbol{\lambda}^{(k)}} l_{comp}$$

in order to remove the unobserved data. The M-step then involves maximizing $Q = \mathbb{E}l_{comp}$ with respect to the parameters in the model to obtain updated estimates of the parameters,

$$(\boldsymbol{\lambda}^{(k+1)}, \boldsymbol{\gamma}^{(k+1)}) = \underset{(\boldsymbol{\lambda},\boldsymbol{\gamma})}{\arg\max}\, Q^{(k)}(\boldsymbol{\lambda}, \boldsymbol{\gamma}).$$

These steps are then alternated until a convergence criterion monitoring subsequent differences in the parameter estimates/likelihood is met.

The calculation of $Q^{(k)}(\boldsymbol{\lambda}, \boldsymbol{\gamma})$ is not difficult and is given by:

$$\begin{aligned} Q^{(k)}(\boldsymbol{\lambda}, \boldsymbol{\gamma}) &= \mathbb{E}\log f_{X,Z|\boldsymbol{\lambda}}(X_{jst}, z_{jst}) + \log f_{Y|\boldsymbol{\gamma}}(y_{ist}) \\ &= \sum_{s=1}^{S}\sum_{t=1}^{T}\sum_{j=1}^{J_t} \mathbb{E}\log f_{X,Z|\boldsymbol{\lambda}}(X_{jst}, z_{jst}) + \sum_{s=1}^{S}\sum_{t=1}^{T}\sum_{i=1}^{I_t} \log f_{Y|\boldsymbol{\gamma}}(y) \\ &\propto \sum_{s,t,j} \left(-\lambda_{st} + z_{jst}\log\lambda_{st}\mathbb{E}X_{jst}\right) + \sum_{s,t}\left(-I_t\gamma_{st} + Y_{\cdot st}\log I_t\gamma_{st}\right) \\ &\propto \sum_{s,t}\left(-J_t\lambda_{st} + z_{\cdot st}\log\lambda_{st}\mathbb{E}(X_{jst}|\lambda_{st}^{(k)}, \gamma_{st}^{(k)})\right) + \sum_{s,t}\left(-I_t\gamma_{st} + Y_{\cdot st}\log I_t\gamma_{st}\right). \end{aligned} \tag{2}$$

No analytic solution to the M-step exists, however, so we chose to maximize Q with coordinate descent (Luo and Tseng, 1992). In fact, as we only need to find parameters that increase the value of Q on each iteration, we forgo fully iterating to find the maximum and instead perform just one step uphill within each EM iteration. Since $Q^{(k)}$ is concave and smooth in the parameters $\boldsymbol{\lambda}, \boldsymbol{\gamma}$, we are able to use the convergence of parameter estimates, $||(\boldsymbol{\lambda}^{(k)}, \boldsymbol{\gamma}^{(k)}) - (\boldsymbol{\lambda}^{(k+1)}, \boldsymbol{\gamma}^{(k+1)})||_\infty < \tau$, for some $\tau > 0$, as our stopping criterion.

As we show in our simulation study, this generalized EM algorithm accurately estimates the parameters when values of λ_{st} are relatively small, such that zeros are prevalent in the data Z_{jst}. In this case, not too much information is lost since estimation of $\mathbb{E}Z_{jst}$ can be estimated well by the proportion of observed zeros. In contrast, if the predator consistently eats a given prey species, few to no zeros will show up in the observed data and $\mathbb{E}Z_{jst}$ is estimated to be nearly 1. The loss of information is best seen by attempting to solve for λ_{st} in the equation $1 = \mathbb{E}Z_{jst} = 1 - \exp\{-\lambda_{st}\}$. As the proportion of ones in the observed data increases, we expect λ_{st} to grow exponentially large. When no zeros are present in the data, so that where only ones are observed, the likelihood can be made arbitrarily large by sending the parameter off to infinity.

2.4 Testing

The likelihood ratio test statistic is

$$\Lambda(X,Y) := -2\log\frac{\sup_{\theta_0} L(\theta_0|X,Y)}{\sup_{\theta_1} L(\theta_1|X,Y)},$$

where θ_0, θ_1 represent the parameters estimated under the null and alternative hypotheses, respectively. It is well known that the asymptotic distribution of Λ is a χ^2_ρ distribution with ρ degrees of freedom (Wilks, 1938). When the observations X_{jst} are not observed, we use $L_{obs}(Z,Y)$ as the likelihood in the calculation of Λ.

The degrees of freedom ρ equal the number of free parameters available in the stated hypotheses under question. If we put the null hypothesis to be $H_0 : \lambda_t = c_t\gamma_t$, for all t and contrast this against $H_1 : \lambda_{st} = c_{st}\gamma_{st}$ then there are $\rho = 2(S\cdot T) - S\cdot T - T = S\cdot T - T$ degrees of freedom.

A set of hypotheses is determined by the p-value of the χ^2_ρ distribution. Hence, with a level of significance, α, the null hypothesis is rejected in favor of the alternative hypothesis if $\mathbb{P}(\chi^2_\rho > \Lambda) < \alpha$.

2.5 Linear Transformations of c_{st}

After determining which model best fits the data, more detail can be extracted through a hypothesis test of the elements of c_{st}, or in vector notation as $\mathbf{c} \in \mathbb{R}^{S\cdot T}$. Let the elements of $\hat{\mathbf{c}}$ be the maximum likelihood estimates, $\hat{c}_{st}$, as found via the framework above. Since $\hat{\mathbf{c}}$ is asymptotically normally distributed, any linear combination of the elements is also asymptotically normally distributed. For instance, let a be a vector of the same dimension of $\hat{\mathbf{c}}$. Then $a^t\hat{\mathbf{c}}$ is asymptotically distributed as $\mathcal{N}(a^t\mathbf{c}, a^t\Sigma a)$, where Σ is the covariance matrix of the asymptotic distribution of $\hat{\mathbf{c}}$.

Suppose, for example, that the hypothesis c_s is determined to best fit the data with s ranging $s = 1, 2, 3$. We can test to see whether or not two species are statistically equally preferred under the null hypothesis $c_1 = c_2$. This hypothesis is alternatively written in vector notation as $a^t\mathbf{c} = 0$, where $a = (1, -1, 0)^t$. Tests of the following form $H_0 : a^t\mathbf{c} = \mu$ against any alternative of interest are then approximate Z-tests. Confidence intervals of any size are similarly, readily obtained. Of course, one could also use a t distribution as a small sample size correction.

3 Simulation Study

Our simulations assume two prey species and five time points, throughout. Of the hierarchy of hypotheses, we generate data under three models: c, c_s, c_t. Sample sizes for both prey species and predator gut count observations are randomly chosen from four overlapping levels: "small" sample sizes are randomly sampled numbers in $[20, 50]$, "medium" $[30, 75]$, "large" $[50, 150]$, and "larger" $[100, 200]$. This is repeated for each level of sample size. We simulate 500 replicate datasets for each of the twelve scenarios above for both types of data, fully observed count data, X_{jst}, and for non-count data, when we observe only a

binary response, $Z_{jst} = 1(X_{jst} > 0)$. Each scenario is then fitted with the true model that generated the data. All simulations of non-count data use $\tau = 10^{-5}$ as the convergence tolerance. A subset of the examples are provided here; the interested reader is referred to the supplementary materials for the complete simulation results. For the simulations we used the `R` Core Team (2014) package `BatchExperiments` by Bischl et al. (2014).

[Figure 2 about here.]

For all simulated data, the true parameter values for the rate at which prey species are encountered in the wild are fixed to be $\gamma_{st} = \pi$, $\forall s, t$. The values of λ_{st} are set with respect to each data generating model. For model $c_{st} = c$, where predator preferences don't vary by either time or species, we put $\lambda_{st} = 2\pi, \forall s, t$. Under model c_s, the ratio of rates vary by species only, so we put $\lambda_{1t} = \sqrt{2}$ and $\lambda_{2t} = \pi$. Hence, $c_1 = \sqrt{2}/\pi \approx 0.45$ and $c_2 = 1$. For the last model, c_t, the ratio of rates vary by time t. Here, we put $\lambda_{st} = t$ for $t \in \{1, \ldots, 5\}$.

[Figure 3 about here.]

Figure 2 shows the density plot of the estimates of c_s when fitting the true model to the fully observed count data generated under models c_s, while figure 3 shows the same for the estimates of c. The plots provide evaluations of parameter estimates under each scenario. For model c_s in figure 2, the parameters $c_1 \approx 0.45$ and $c_2 = 1$ are on average, across all 500 simulations, estimated as $\hat{c}_1 = 0.45$ and $\hat{c}_2 = 1.00$, with standard errors of $\text{se}(\hat{c}_1) = 0.03$ and $\text{se}(\hat{c}_2) = 0.06$. Figure 3 provides results for model $c_{st} = c$. Averaging across all 500 simulations, the parameter $c = 2$ is estimated as $\hat{c} = 2.00$. This is further seen in figure 4, where box plots of the parameter estimates, centered at true parameter values, of the correct model fit to data generated from both c_s and c_t show empirically very little bias.

[Figure 4 about here.]

We next generated data with unobserved counts. As noted above under certain circumstances our unobserved counts model accurately estimates the parameters of interest, and at other times can infinitely over-estimate parameters. To investigate this issue further, we consider the same scenarios mentioned above, but now we reduce all of the count data down to binary observations. For each scenario, we fit the unobserved counts model as if we knew the true underlying model that generated the observed data.

Figures 5 and 6 contain density plots for all 500 replications of the data generating models c_s, c_t with the small and the larger sample sizes, respectively. When data are generated under the model c_s and the true model fit to the non-count data, we find even for the small sample size that point estimates are only very slightly biased. When parameter values are of sufficient size to make zeros in the simulated data less common, the estimates from fitting the correct model to the generated data are occasionally over-estimated. This effect is easily seen in figure 6 for the two greatest values of c_t despite the increased sample size, but is also seen, less dramatically, in the density plot for the c_s generated data.

[Figure 5 about here.]

The cluster of estimates for c_5 between 3.5 and 4.0 in figure 6 comes from datasets in which $X_{js5} > 0$ so that $Z_{js5} = 1$ for all j, s. For the data shown in figure 6, this happened 73 times out of the 500 replicated datasets. As mentioned above, the estimate of c_5 is infinite in this case. However, the EM algorithm will always provide a finite estimate for all parameters when it terminates. In this case, we set $\tau = 10^{-5}$ and this caused the algorithm to terminate with $\hat{c}_5$ between 3.5 and 4.0. To confirm that this is due to the arbitrary choice of τ, we repeated the algorithm with smaller values of τ for several datasets. As expected, $\hat{c}_5$ increased without bound as we refit the model with increasingly small values of τ.

[Figure 6 about here.]

The over-estimation of parameters, a symptom of the loss of information due to the unobserved counts, can also be seen with box plots of the 500 point estimates centered at their respective true parameter values. Figure 7 contains box plots of the same scenarios in figures 5 and 6. For the 73 cases in which $Z_{js5} = 1$ for all j, s under model c_t with the larger sample size, the bias is infinite since parameter estimates will, theoretically, be infinite. The finite bias shown in these plots is due to the finite estimates provided by the termination of the EM algorithm. Thus, conditional on a mixture of 0s and 1s in the data the corresponding estimators are unbiased, but when no 0s exist in the data the theoretical bias is infinite.

[Figure 7 about here.]

4 Application

We analyzed a dataset that was collected to investigate the feeding preferences of two species of wolf spider, Schizocosa ocreata and Schizocosa stridulans (Araneae: Lycosidae). Every $6 - 12$ days, 10 to 40 spiders were hand-collected between October 2011 and April 2013 within Berea College Forest in Madison County, Kentucky, USA. Spiders were removed from the litter using an aspirator, placed in separate 1.5 mL microcentrifuge tubes filled with 95% EtOH, and preserved at -20℃ until DNA extraction. In parallel, we also surveyed availability of forest floor prey using pitfall traps ($n = 32$). For the analysis, both species of Schizocosa were pooled and the number of spiders and prey were analysed by month. On average, 69 spiders, 111 Diptera, and 297 Collembola were caught in each time period. The range of the sample sizes across all 18 months was 11 to 181 for caught spiders, 7 to 322 for trapped Diptera, and 101 to 755 for trapped Collembola. Figure 8 plots the total number of each order that was caught during each time period.

[Table 1 about here.]

To determine whether spiders consumed dipterans and/or collemblans, we conducted a molecular analysis of their gut-contents. First, DNA from spiders was extracted using Qiagen DNEasy®Tissue Extraction Kit (Qiagen Inc., Chatsworth, California, USA) following the animal tissue protocol outlined by the manufacturer, with minor modifications. Whole bodies of the spiders were first crushed to release prey DNA from within their alimentary canal for extraction. The 200μL extractions were stored at -20℃ until PCR. Second, order-specific

primers from the literature were used to detect the DNA of Collembola and Diptera within the guts of the spiders. Primer pairs designed by Sint et al. (2012), targeting the 18S rDNA gene, were used to detect Collembola predation table 7. A PCR cycling protocol for 12.5μL reactions containing $1x$ Takara buffer (Takara Bio Inc., Shiga, Japan), 0.2 mM dNTPs, 0.2μM of each primer, 0.625 U Takara Ex TaqTM and 1.5μL of template DNA, using BioRad PTC$-$200 and C1000 thermal cyclers (Bio-Rad Laboratories, Hercules, California, USA), was optimized as follows: 95℃ for 1 minute, followed by 35 cycles of 94℃ for 30 seconds, 61.2℃ for 90 seconds, and 72℃ for 60 seconds. Primer pairs designed by Eitzinger et al. (2013), targeting the 18S rDNA gene, were used to detect Diptera predation table 7. PCR cycling protocol for 12.5μL reactions with Takara reagents (as above) and 2μL of template DNA was optimized as follows: 95℃ for 1 minute, followed by 40 cycles of 94 for 45 seconds, 60℃ for 45 seconds, and 72℃ for 45 seconds. Both primer pairs were tested for cross-reactivity against a range of prey and predator species from the field site and in all cases, no amplification of DNA was observed, confirming suitable specificity of the primers for this study. Lastly, electrophoresis of 10μL of each PCR product was later conducted to determine success of DNA amplification using 2% Seakem agarose (Lonza, Rockland, Maine, USA) stained with $1x$ GelRed™nucleic acid stain (Biotium, Hayward, California, USA). This procedure allowed us to determine a presence or an absence of Diptera and Collembola DNA within each spider.

[Figure 8 about here.]

[Figure 9 about here.]

These data provide an example of our hierarchy of hypotheses. First, we tested model $c_{st} = c$ against $c_{st} = c_s$, to determine whether or not the wolf spider has different preferences for the two orders Diptera and Collembola. With, one degree of freedom, this likelihood ratio test indicated, $p - value < 0.0001$, that two parameters, one for each order, fits these data better than one parameter for both. Similarly, we tested whether or not there was a significant effect across time by testing model $c_{st} = c$ against $c_{st} = c_t$. Here, the likelihood ratio test, with 17 degrees of freedom, implies that the wolf spiders of the Berea College Forest eat these prey orders at different rates across the months of the year, $p - value < 0.0001$. In fact, we find that the most parameter rich model, $\lambda_{st} = c_{st}\gamma_{st}$ fits these data better than is expected by chance, $p - value < 0.0001$. Model c_{st} estimates 72 parameters in total; since, in this case, there are two prey of interest and 18 time periods, it takes 36 parameters to estimate each c_{st} and γ_{st}. Figures 10, 11 plot the point estimates and 95% confidence intervals of c_{st}, for both prey across all time periods.

[Figure 10 about here.]

[Figure 11 about here.]

With point estimates of c_{st} under the model $\lambda_{st} = c_{st}\gamma_{st}$, we can test any number of linear contrasts. For instance, $c_{1t} = c_{2t}$, for $t \in \{1, \ldots, 18\}$. Using a level of significance of 0.05, and after making a Bonferroni multiple comparisons adjustment, the data can not say that the two prey are differently preferred in October, November, and December of 2011 and for March and July of 2012.

5 Discussion

The model developed here allows for the determination of a predator's preferences by testing simultaneously across an array of multiple prey species and time points. This is achieved via a simple, but statistically powerful, likelihood ratio test. Further testing of the ratio of rates for which predators eat to encounter prey species allows researchers to make specific conclusions about predators' preferences. For instance, rates across time can be estimated to make statements about seasonal effects on a predator's eating habits, or relative rates across species groups allows for statements about the relative preferences for different species.

When counts of predators' gut contents are not fully observed, and instead only a binary response indicating the existence of the prey species in the gut is observed, we are able to treat the counts as missing data. By modeling all of the observed data, both the binary responses and the number of prey species caught, and the missing count data, we are able to use the EM algorithm to extract as much information from the data as possible. Though this is nice in theory, in practice the success of this modification to our original model is limited by the magnitude of the unknown parameters λ_{st}.

Further developments of our model could be beneficial. Taking into account other environmental variables that might effect a predator's eating habits, such as rain or temperature might also be advantageous.

6 Acknowledgments

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7 Data Accessibility

An `R` package, named `spiders`, is available on CRAN at `http://cran.r-project.org/web/packages/spiders/index.html` and fits all the methods discussed above.

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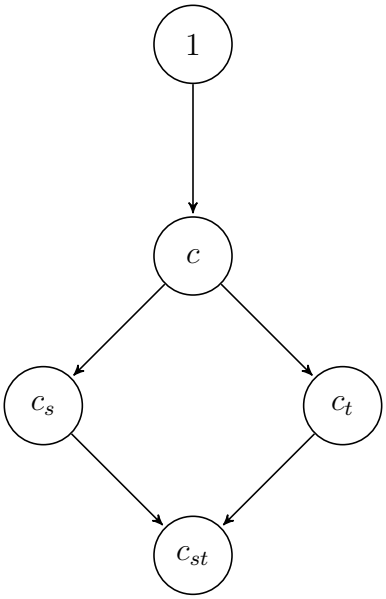

Figure 1: This hierarchy of hypotheses suggests the order in which the discussed models should be tested. One begins with the models at the top and sequentially, following the arrows, tests the hypotheses using the formal test described in section 2.4.

Small Sample Size

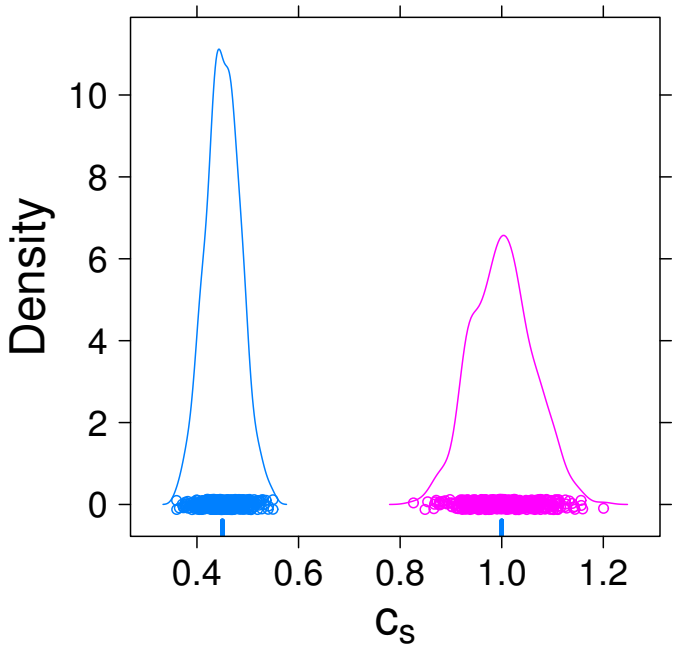

Figure 2: The density plot of all 500 estimates of fitting the true model to the data generated from models c_s with the small sample sizes is shown. Each element of c_s is color coded for clarity, and ticks on the x-axis show the true parameter values.

Medium Sample Size

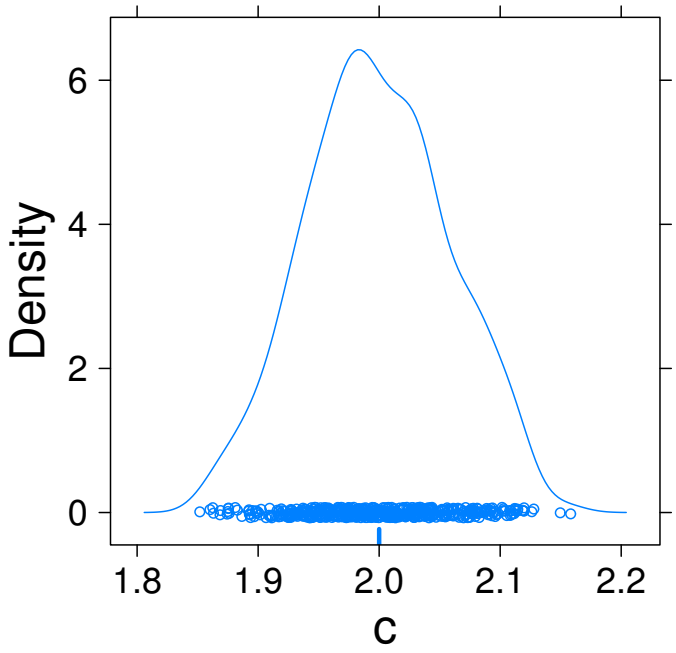

Figure 3: The density plot of all 500 estimates of fitting the true model to the data generated from model c with the small sample sizes is shown. A tick on the x-axis shows the true parameter value.

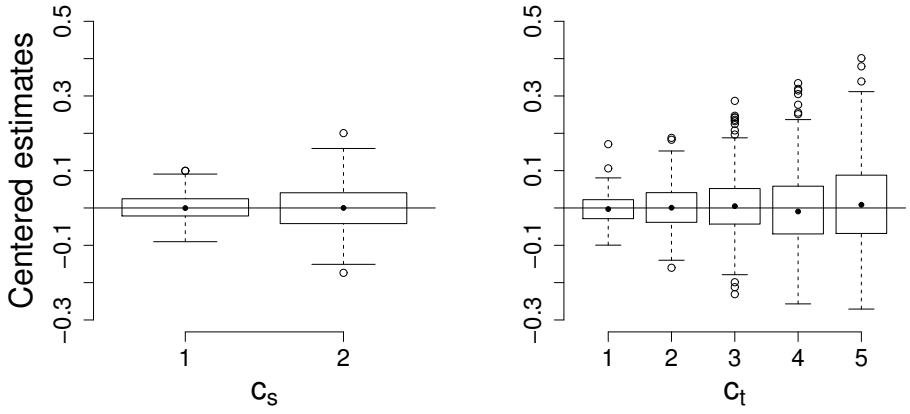

Figure 4: Shown are all 500 estimates, centered at the true parameter values, from fitting the true model to the data generated from models c_s, c_t with sample sizes smallest and small, respectively.

Small Sample Size, Non-Count Data

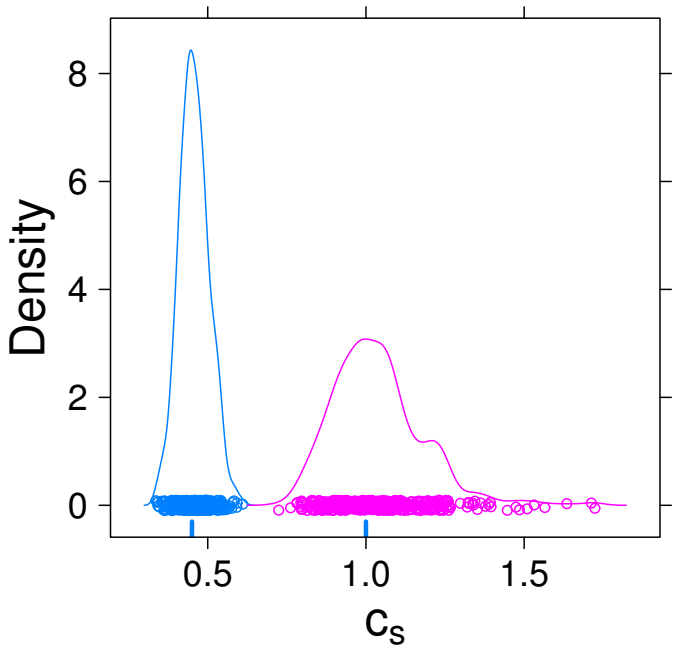

Figure 5: The density plot of all 500 estimates of fitting the true model to the data generated from model c_s, when counts are not observed, is shown with the small sample size. Each element of c_s is color coded for clarity, and ticks on the x-axis show the true parameter values.

Medium Sample Size, Non-Count Data

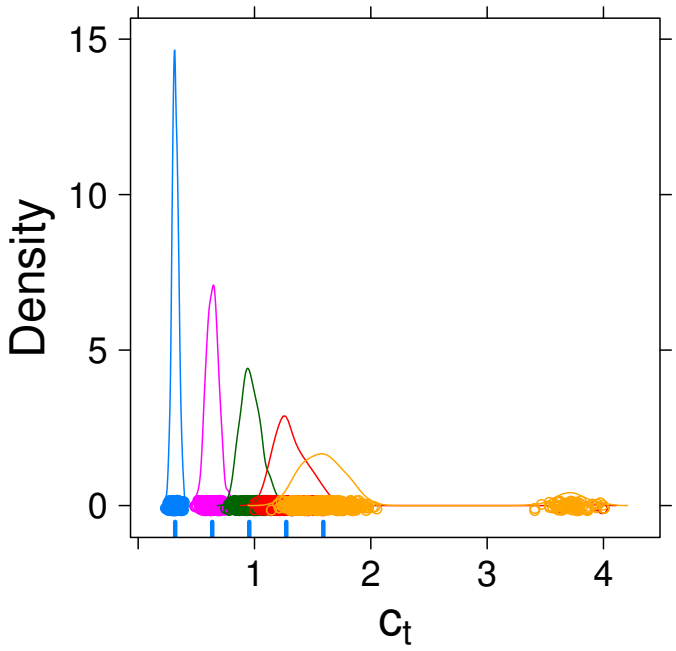

Figure 6: The density plot of all 500 estimates of fitting the true model to the data generated from model c_t, when counts are not observed, is shown with the larger sample size. Each element of c_t is color coded for clarity, and ticks on the x-axis show the true parameter values.

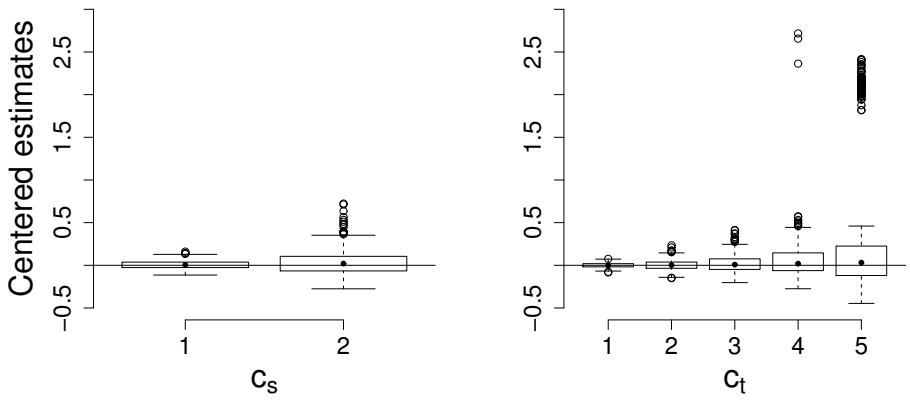

Figure 7: Shown are all 500 estimates, centered at the true parameter values, from fitting the true model to the data generated from models c_s, c_t, when counts are not observed, with sample sizes small and larger, respectively.

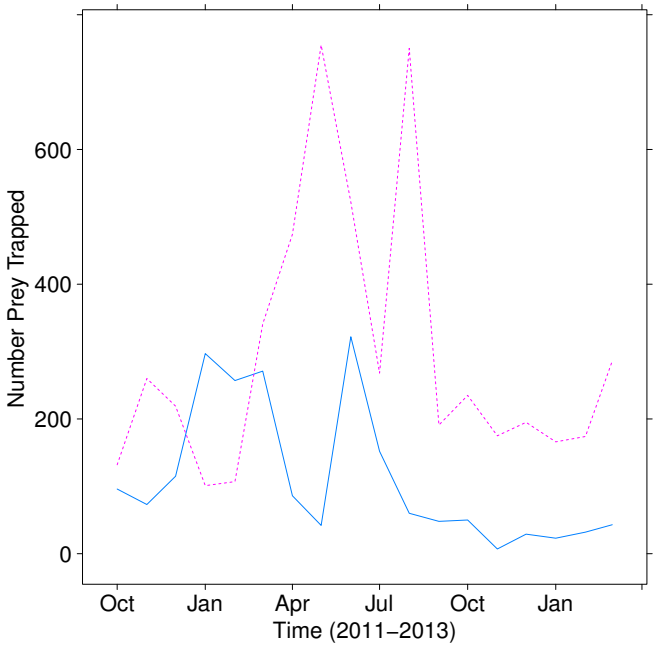

Figure 8: For both Collembola (pink/dashed) and Diptera (blue/solid), the plot shows the number of the prey trapped in each time period.

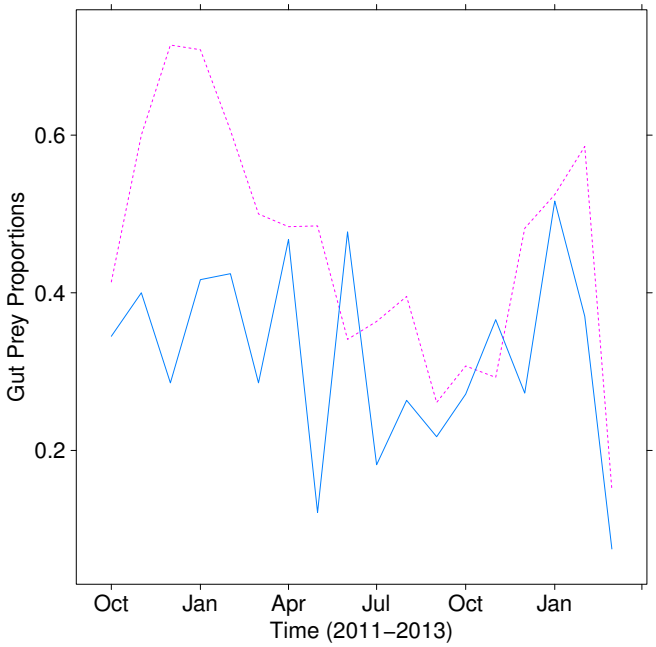

Figure 9: For both Collembola (pink/dashed) and Diptera (blue/solid), the plot shows the prey proportions in the sampled wolf spiders' guts in each time period.

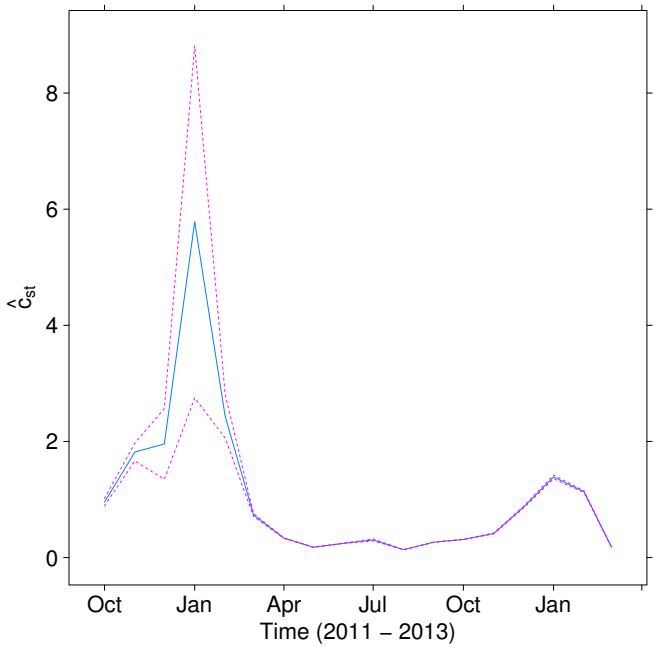

Figure 10: Point estimates (blue/solid) and 95% confidence intervals (pink/dashed) as estimated from the model c_{st} for Collembola.

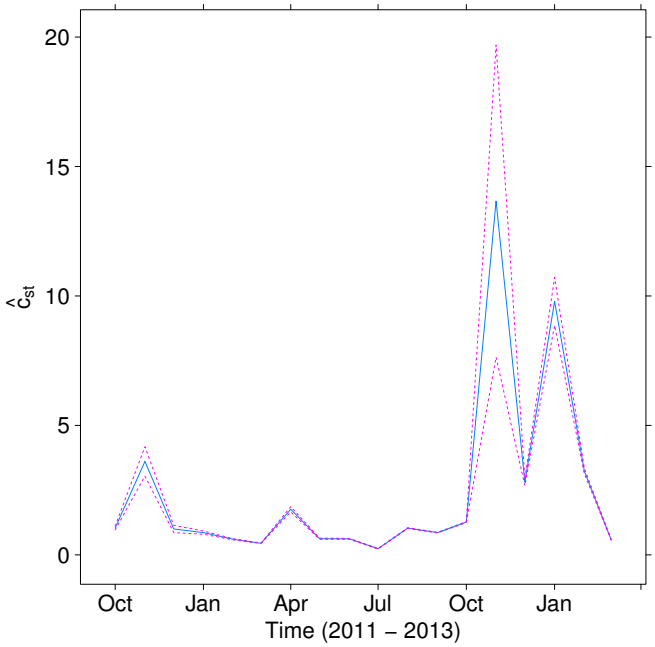

Figure 11: Point estimates (blue/solid) and 95% confidence intervals (pink/dashed) as estimated from the model c_{st} for Diptera.

Target group	**Primer names and sequences** $5' - 3'$	**Size (bp)**	**Source**
Collembola	Col3F: GGACGATYTTRTTRGTTCGT	228	Sint et al. 2012
	Col-gen-A246: TTTCACCTCTAACGTCGCAG		
Diptera	DIPS16: CACTTGCTTCTTAAATrGACAAATT	198	Eitzinger et al. 2013
	DIPA17: TTyATGTGAACAGTTTCAGTyCA		

Table 1: Targeted prey orders, primer names and sequences, size of amplicon, and source of design for the detection of prey taxa within the guts of Schizocosa spiders. Both primer sets were used in singleplex PCR assays.