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POPULATION GENETICS MODELS WITH SKEWED FERTILITIES: A FORWARD AND BACKWARD ANALYSIS

THIERRY HUILLET, MARTIN MÖHLE

ABSTRACT. Discrete population genetics models with unequal (skewed) fertilities are considered, with an emphasis on skewed versions of Cannings models, conditional branching process models in the spirit of Karlin and McGregor, and compound Poisson models. Three particular classes of models with skewed fertilities are investigated, the Wright-Fisher model, the Dirichlet model, and the Kimura model. For each class the asymptotic behavior as the total population size N tends to infinity is investigated for power law fertilities and for geometric fertilities. This class of models can exhibit a rich variety of sub-linear or even constant effective population sizes. Therefore, the models are not necessarily in the domain of attraction of the Kingman coalescent. For a substantial range of the parameters, discrete-time coalescent processes with simultaneous multiple collisions arise in the limit.

Running title: Population genetics models with skewed fertilities.

Keywords: Ancestral process; Cannings model; Compound Poisson model; Dirichlet model; Dirichlet–Kingman coalescent; Duality; Evolutionary processes; Exchangeable coalescent; Karlin and McGregor model; Kimura model; Kingman coalescent; Population dynamics; Simultaneous multiple collisions; Wright-Fisher model

1. INTRODUCTION

The well-known neutral discrete Wright-Fisher model describes the evolution of a population of constant size N from generation to generation forwards in time. One of the main features of this model (and similar models such as the discrete Moran model) is that each individual has the same propensity to produce offspring.

In this paper discrete population models are studied, in which individuals may have unequal propensities to reproduce. We shall speak of models with skewed fertilities. The terminology ‘skewed’ means here ‘asymmetric’ or ‘unequal’ and should be not confused with the skewness of a distribution. These models turn out to be of interest mainly because of the following two reasons. First of all, it turns out that the effective population size N_e (a precise definition of N_e is provided after Eq. (8)) of these models may differ significantly from the actual total population size N and may not even depend linearly on N . The fact that models with this behavior may have importance for biological applications, is for example indicated in papers of Eldon and Wakeley [6] and Wakeley and Sargsyan [34]. Models with effective population sizes smaller than N occur also in a different context of age-structure of populations in a paper of Sagitov and Jagers [30]. Secondly, in the limit as the total population size N tends to infinity, these models are not necessarily in

the domain of attraction (see Definition 2.1) of the Kingman coalescent [14]. The limiting coalescent may have simultaneous multiple collisions of ancestral lineages. Ancestral processes of this form are a major research area in coalescent theory and have been studied intensively since at least ten years (see, for example, [24] or [31]).

The models we are mainly interested in are obtained by conditioning a sequence $\xi_1, \xi_2, \dots$ of independent random variables on the event that $\xi_1 + \dots + \xi_N = N$, with the interpretation that the outcome of ξ_n , under the constraint that $\xi_1 + \dots + \xi_N = N$, is the number of children of the n th individual, $n \in \{1, \dots, N\}$. Since the random variables $\xi_1, \xi_2, \dots$ are not necessarily assumed to be identically distributed, this construction results in unequal propensities of the individuals to produce offspring. In the spirit of Karlin and McGregor [13] we call this model the skewed conditional branching process model.

Schweinsberg [32] studies models, in which N individuals are sampled without replacement from $S_N := X_1 + \dots + X_N$ offspring, where $X_1, \dots, X_N$ are given i.i.d. random variables. The effective population sizes of these models satisfy $N_e = O(N)$ (see [32, Eq. (18)]), and, in particular, N_e can exhibit a rich variety of sub-linear behavior (see, for example, [32, Lemma 13]), however, these models are different from our models, which are based on conditioning instead of sampling.

The paper is organized as follows. In Section 2 a skewed Cannings model is introduced and analyzed. This model essentially coincides with an associated standard exchangeable Cannings model [3, 4]. Section 2 therefore heavily gains from the theory on Cannings models. The results of Section 2 are applied in Section 3 to the skewed conditional branching process model. Since this huge class of conditional models has in general a quite complicated probabilistic structure, we further specialize in Section 4 to a subclass of skewed compound Poisson models. Exact formulae for the transition probabilities of the forward and the backward process are derived. It is furthermore shown (Theorem 4.3) that, in the unbiased (non-skewed) case, the model is in the domain of attraction (see Definition 2.1) of the Kingman coalescent. In all cases we obtain exact and asymptotic formulae for the effective population size N_e , which can deviate substantially from the total population size N . The following Sections 5 and 6 are devoted to two particular compound Poisson models, the skewed Wright-Fisher model and the skewed Dirichlet model. In both models the effective population size N_e is less than or equal to the total population size N , and, depending on the parameter choices, can indeed be substantially smaller than N , for example $N_e \sim \rho N$ with $\rho \in (0, 1)$, $N_e = O(N^\beta)$ with $\beta \in (0, 1)$, $N_e = O(\log N)$, or even $N_e = O(1)$. These two models are in particular analyzed for power law skewed fertilities and for geometrically skewed fertilities. It turns out that for a wide range of parameter choices, these models are not anymore in the domain of attraction of the Kingman coalescent. Coalescents allowing for simultaneous multiple collisions arise in the limit as the total population size tends to infinity. The paper finishes in Section 7 with an analog analysis of the skewed Kimura model, a model which does not belong to the compound Poisson class, but nevertheless exhibits similarities with the previously studied models.

Throughout the paper the notation $\mathbb{N} := \{1, 2, \dots\}$, $\mathbb{N}_0 := \{0, 1, 2, \dots\}$, and $[N] := \{1, \dots, N\}$, $N \in \mathbb{N}$, is used. We furthermore use, for fixed $N \in \mathbb{N}$, the symbol $S := \{0, \dots, N\}$ for the state space of several discrete processes considered in this paper. Moreover, for $k \in \mathbb{N}_0$ and $x \in \mathbb{R}$, $(x)_k := x(x-1)\dots(x-k+1)$ and

$[x]_k := x(x+1)\cdots(x+k-1)$ denote the descending and ascending factorials respectively, with the convention that $(x)_0 = [x]_0 = 1$.

2. SKEWED CANNINGS MODELS

Consider a population with constant population size $N \in \mathbb{N}$ evolving in discrete non-overlapping generations $t \in \mathbb{N}_0$. The n th individual, $n \in [N]$, of the t th generation gives birth to a random number $\mu_n(t)$ of children. Since the population is assumed to have constant population size N , for each fixed generation $t \in \mathbb{N}$, the random vector $\mu(t) := (\mu_1(t), \dots, \mu_N(t))$ must take values in the discrete N -simplex $\Delta(N)$ consisting of all $k = (k_1, \dots, k_N) \in \mathbb{N}_0^N$ satisfying $k_1 + \dots + k_N = N$. Note that the random variable $\mu_n(t) = \mu_{n,N}(t)$ is allowed to depend on the total population size N . However, for simplicity, this dependence on N is usually not indicated throughout the paper. It is assumed that the random vectors $\mu(0), \mu(1), \dots$ are i.i.d., so the model is time-homogeneous. We write $\mu_n := \mu_n(0)$, $n \in [N]$, and $\mu := \mu(0)$ for convenience. Note that the model is in principle defined in the same way as an exchangeable Cannings model [3, 4]. The only slight modification is that, for each fixed generation $t \in \mathbb{N}_0$, the offspring variables $\mu_1(t), \dots, \mu_N(t)$ are not necessarily assumed to be exchangeable. Models of this form have been studied in [19] even in the time-inhomogeneous setting when the total population size is not necessarily constant equal to N . Particular examples, such as the skewed Wright-Fisher model, which is studied in Section 5 in more detail, have been the source of recurrent interest in the literature (see, for example, [2] and [18]). A main tool exploited in this section is the following shuffling procedure. For each fixed generation $t \in \mathbb{N}_0$ let $\nu(t) = (\nu_1(t), \dots, \nu_N(t))$ be a random permutation (shuffling) of $\mu(t) = (\mu_1(t), \dots, \mu_N(t))$. Then $\nu(t)$ is exchangeable with distribution

$$(1) \quad \mathbb{P}(\nu(t) = k) = \frac{1}{N!} \sum_{\pi} \mathbb{P}(\pi\mu(t) = k), \quad k \in \Delta(N),$$

where $\pi\mu(t) := (\mu_{\pi_1}(t), \dots, \mu_{\pi_N}(t))$ and the sum extends over all permutations π of $[N]$. We interpret $\nu_n(t)$ as the number of offspring of the n th individual of the t th generation in an exchangeable Cannings model [3, 4]. Again we write $\nu_n := \nu_n(0)$, $n \in [N]$, and $\nu := \nu(0)$ for convenience. As we will see soon, this exchangeable Cannings population model will turn out to be very helpful to analyze the original skewed Cannings model. More precisely, since all assignments of offspring to parents are assumed to be equally likely, there is essentially no difference between the skewed model and the associated shuffled exchangeable Cannings model. However, we are interested in these skewed Cannings models since (as the following sections will show) starting with not necessarily exchangeable offspring variables $\mu_1, \dots, \mu_N$ naturally leads to interesting subclasses and examples of Cannings models, which to the best of our knowledge have not been introduced so far. Most of the results provided in this section are well known from the literature on Cannings models and from coalescent theory. We chose to present these results, since we interpret them in terms of the original skewed model we started with at the beginning of this section. For example, we clarify how the forward and backward processes of these models depend on the original offspring variables $\mu_1, \dots, \mu_N$.

2.1. Looking forwards in time. Take a sample of $n \in S$ individuals of generation 0 and, for $t \in \mathbb{N}_0$, let X_t denote the number of descendants of these n individuals

in generation t . The process $X := (X_t)_{t \in \mathbb{N}_0}$, called the forward process, is a time-homogeneous Markov chain with state space S and initial state $X_0 = n$. The transition probabilities $P_{i,j} := \mathbb{P}(X_{t+1} = j \mid X_t = i)$, $i, j \in S$, must coincide with those of the associated exchangeable Cannings model with offspring distributions (1), since in the considered skewed model all assignments of offspring to parents are assumed to be equally likely. From the literature on exchangeable models (Cannings [3, p. 267]), it follows that $P_{i,j} = \mathbb{P}(\nu_1 + \dots + \nu_i = j) = \sum_k \mathbb{P}(\nu = k)$, where the sum extends over all $k \in \Delta(N)$ satisfying $k_1 + \dots + k_i = j$. Plugging in (1) and interchanging the two sums involved yields

$$P_{i,j} = \frac{1}{N!} \sum_{\pi} \sum_k \mathbb{P}(\pi\mu = k) = \frac{1}{N!} \sum_{\pi} \mathbb{P}(\mu_{\pi 1} + \dots + \mu_{\pi i} = j).$$

Introducing $n_1 := \pi 1, \dots, n_i := \pi i$, and noting that there exist exactly $(N-i)!$ permutations π leaving $\pi 1, \dots, \pi i$ fixed, it follows that

$$(2) \quad P_{i,j} = \frac{1}{(N)_i} \sum_{\substack{n_1, \dots, n_i=1 \\ \text{all distinct}}}^N \mathbb{P}\left(\sum_{k=1}^i \mu_{n_k} = j\right) = \frac{1}{\binom{N}{i}} \sum_{1 \leq n_1 < \dots < n_i \leq N} \mathbb{P}\left(\sum_{k=1}^i \mu_{n_k} = j\right)$$

for $i, j \in S$. One may rewrite (2) as $P_{i,j} = \binom{N}{i}^{-1} \sum_M \mathbb{P}(\sum_{m \in M} \mu_m = j)$, $i, j \in S$, where the sum $\sum_M$ extends over all subsets M of $[N]$ satisfying $|M| = i$. The chain X is a martingale, since $\mathbb{E}(X_{t+1} \mid X_t = i) = \sum_{j \in S} j P_{i,j} = \sum_{j \in S} j \mathbb{P}(\nu_1 + \dots + \nu_i = j) = \mathbb{E}(\nu_1 + \dots + \nu_i) = i \mathbb{E}(\nu_1) = i$, $i \in S$. Since X is bounded ($0 \leq X_t \leq N$ for all $t \in \mathbb{N}_0$), X_t converges almost surely to some random variable X_{∞} as $t \rightarrow \infty$, and $(X_t)_{t \in \mathbb{N}_0 \cup \{\infty\}}$ is still a martingale. If $\mathbb{P}(\mu_k = 1) < 1$ for some $k \in [N]$, then (see, for example, Section 2.1 of [20]) X_{∞} takes the two values 0 and N with probability $\mathbb{P}(X_{\infty} = 0) = 1 - n/N$ and $\mathbb{P}(X_{\infty} = N) = n/N$ respectively.

The associated Cannings model with exchangeable offspring variables $\nu_1, \dots, \nu_N$ is useful in many respects. For instance, in terms of the so-called coalescence probability (see, for example, [21])

$$(3) \quad c_N := \frac{\mathbb{E}((\nu_1)_2)}{N-1} = \frac{\text{Var}(\nu_1)}{N-1} = 1 - \mathbb{E}(\nu_1 \nu_2) = -\text{Cov}(\nu_1, \nu_2),$$

the variance of X_{t+1} , given $X_t = i$, can be expressed as

$$\begin{aligned} \text{Var}(X_{t+1} \mid X_t = i) &= \text{Var}(\nu_1 + \dots + \nu_i) = i \text{Var}(\nu_1) + i(i-1) \text{Cov}(\nu_1, \nu_2) \\ &= i(N-1)c_N - i(i-1)c_N = i(N-i)c_N. \end{aligned}$$

Defining the heterogeneity of the population at generation t as $2(X_t/N)(1 - X_t/N)$, we have $2\mathbb{E}((X_t/N)(1 - X_t/N) \mid X_0 = i) = 2(1 - c_N)^t (i/N)(1 - i/N)$, $t \in \mathbb{N}_0, i \in S$, showing that, if $c_N > 0$, the mean heterogeneity tends to 0 exponentially fast as $t \rightarrow \infty$. We will provide further information on c_N when the model is studied backwards in time.

2.2. Looking backwards in time. Instead of looking forwards in time let us now look backwards in time and count, starting with all N individuals from some generation $t_0 \in \mathbb{N}_0$, the number of ancestors of these N individuals $t \in \{0, \dots, t_0\}$ generations backward in time. More precisely, let $\hat{X}_t$ denote the number of ancestors of the N individuals of generation t_0 in generation $t_0 - t$. The process $\hat{X} := (\hat{X}_t)_{t \in \{0, \dots, t_0\}}$, called the backward process, is a time-homogeneous Markov chain with state space

S and initial state N . The transition probabilities $\hat{P}_{i,j} := \mathbb{P}(\hat{X}_{t+1} = j \mid \hat{X}_t = i)$, $i, j \in S$, coincide with those of the associated Cannings model with exchangeable offspring vector ν having distribution (1). Therefore, from the literature on exchangeable models (Cannings [3, Theorem 11], Gladstien [9, Examples]), it follows that

$$\begin{aligned}
 \hat{P}_{i,j} &= \frac{\binom{N}{j}}{\binom{N}{i}} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \mathbb{E} \left(\prod_{k=1}^j \binom{\nu_k}{l_k} \right) \\
 (4) \quad &= \frac{1}{\binom{N}{i}} \sum_{1 \leq n_1 < \dots < n_j \leq N} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \mathbb{E} \left(\prod_{k=1}^j \binom{\mu_{n_k}}{l_k} \right), \quad i, j \in S,
 \end{aligned}$$

with the convention that $\hat{P}_{i,0} = \delta_{i0}$ (Kronecker symbol). Note that the last equality follows from (1) via a similar argument as already used to derive the formula (2) for $P_{i,j}$. The states 0 and 1 are absorbing.

The matrix $\hat{P} = (\hat{P}_{i,j})_{i,j \in S}$ is lower left triangular and has hence eigenvalues $\lambda_i := \hat{P}_{i,i}$, $i \in S$. From (4) it follows that $\lambda_0 = 1$ and that $\lambda_i = \hat{P}_{i,i} = \binom{N}{i}^{-1} \sum_{1 \leq n_1 < \dots < n_i \leq N} \mathbb{E}(\mu_{n_1} \cdots \mu_{n_i})$, $i \in [N]$. Note that $\lambda_1 = 1$ and that $\lambda_2 = \binom{N}{2}^{-1} \sum_{1 \leq n_1 < n_2 \leq N} \mathbb{E}(\mu_{n_1} \mu_{n_2}) = \mathbb{E}(\nu_1 \nu_2) = 1 - c_N$. Since the forward and backward transition matrices $P = (P_{i,j})_{i,j \in S}$ and $\hat{P} = (\hat{P}_{i,j})_{i,j \in S}$ coincide with those of the associated Cannings model with exchangeable offspring vector ν , it is allowed to apply results on exchangeable Cannings models, in particular duality results such as the duality relation $PH = H\hat{P}'$, where $H = (H_{ij})_{i,j \in S}$ is (see, for example, [20]) the matrix with entries $H_{ij} = \binom{i}{j} / \binom{N}{j}$, $i, j \in S$. Since H is non-singular, P has the same eigenvalues as $\hat{P}$. In particular, $c_N = 1 - \lambda_2$ is the spectral gap of P and $\hat{P}$.

The random variable $\hat{X}_t$ counts the number of ancestors in generation $t_0 - t$, but it gives no information about whether two individuals i and j , randomly picked from generation t_0 , share a common parent in generation $t_0 - t$. In order to encode this information, a more enriched ancestral process has to be considered, which is now described. Take a random sample of $n \in [N]$ individuals from generation $t_0 \in \mathbb{N}_0$ and, for $t \in \{0, \dots, t_0\}$, define a random relation $\mathcal{R}_t$ on $\{1, \dots, n\}$ by saying that $(i, j) \in \mathcal{R}_t$ if and only if the individuals i and j have a common parent in generation $t_0 - t$. Note that $\mathcal{R}_t = {}_{(N)}\mathcal{R}_t^{(n)}$ depends on the sample size n and on the total population size N . The process $(\mathcal{R}_t)_{t \in \{0, \dots, t_0\}}$ is called the ancestral process or a discrete coalescent process. It is well known (see, for example, [19]) that $(\mathcal{R}_t)_{t \in \{0, \dots, t_0\}}$ is a Markov chain with state space $\mathcal{E}_n$, the set of equivalence relations (partitions) on $\{1, \dots, n\}$, and transition probabilities

$$(5) \quad \mathbb{P}(\mathcal{R}_{t+1} = \eta \mid \mathcal{R}_t = \xi) = \Phi_j(l_1, \dots, l_j), \quad \xi, \eta \in \mathcal{E}_n, \xi \subseteq \eta,$$

where

$$(6) \quad \Phi_j(l_1, \dots, l_j) := \frac{1}{\binom{N}{l_1 + \dots + l_j}} \sum_{\substack{n_1, \dots, n_j = 1 \\ \text{all distinct}}}^N \mathbb{E}((\mu_{n_1})_{l_1} \cdots (\mu_{n_j})_{l_j}).$$

Here $j := |\eta|$ denotes the number of equivalence classes (blocks) of η and $l_1, \dots, l_j \in \mathbb{N}$ are the group sizes of merging classes of ξ . Note that $l_1 + \dots + l_j$ is the number of classes (blocks) of ξ . Comparing (6) with (4) shows that

$$(7) \quad \hat{P}_{i,j} = \frac{i!}{j!} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{\Phi_j(l_1, \dots, l_j)}{l_1! \dots l_j!}, \quad i, j \in \{1, \dots, N\}.$$

For $j = 1$, (4) and (7) reduce to

$$\hat{P}_{i,1} = \Phi_1(i) = \frac{1}{(N)_i} \sum_{n=1}^N \mathbb{E}((\mu_n)_i), \quad i \in \{1, \dots, N\}.$$

The coalescence probability c_N , i.e. the probability that two individuals, randomly chosen from some generation, have a common parent, is hence

$$(8) \quad c_N := \hat{P}_{2,1} = \Phi_1(2) = \frac{1}{(N)_2} \sum_{n=1}^N \mathbb{E}((\mu_n)_2),$$

in agreement with (3), and the effective population size is $N_e := 1/c_N$. We will later also make use of the probability that three individuals, randomly chosen from some generation, share a common parent, which is given by

$$d_N := \hat{P}_{3,1} = \Phi_1(3) = \frac{1}{(N)_3} \sum_{n=1}^N \mathbb{E}((\mu_n)_3).$$

The transition probabilities (5) do not depend on t and t_0 . It is hence allowed to choose t_0 arbitrary large. We can therefore think of a process $(\mathcal{R}_t)_{t \in \mathbb{N}_0}$ with transition probabilities (5) and time $t \in \mathbb{N}_0$. A fundamental result from coalescent theory (see, for example, [21, p. 989] or [25, Lemma 5.5]) states that, if $d_N/c_N \rightarrow 0$, then $c_N \rightarrow 0$ and, moreover (see [24] or [21, Theorem 4 (b)]), if $d_N/c_N \rightarrow 0$, then, for each sample size $n \in \mathbb{N}$, the time-scaled process $(\mathcal{R}_{[t/c_N]}^{(n)})_{t \in [0, \infty)}$ converges weakly to Kingman's n -coalescent $(R_t^{(n)})_{t \in [0, \infty)}$ as the total population size N tends to infinity. In [24] there is also a criterion in terms of the quantities (6) provided ensuring that, for each $n \in \mathbb{N}$, the time-scaled process $(\mathcal{R}_{[t/c_N]}^{(n)})_{t \in [0, \infty)}$ converges weakly as $N \rightarrow \infty$ to a more general process $(\varrho_n R_t)_{t \in [0, \infty)}$, where ϱ_n denotes the restriction from $\mathcal{E}$, the set of all equivalence relations on $\mathbb{N}$, to $\mathcal{E}_n$, and $R = (R_t)_{t \geq 0}$ is a continuous-time coalescent process allowing for simultaneous multiple collisions of ancestral lineages. This asymptotic behavior can only occur if $c_N \rightarrow 0$. If, instead, c_N converges to a positive constant, then, a similar criterion for the quantities (6), also provided in [24], ensures that, for each sample size $n \in \mathbb{N}$, the process $(\mathcal{R}_t^{(n)})_{t \in \mathbb{N}_0}$ (without any time-scaling involved) converges weakly to a discrete-time process $(\varrho_n R_t)_{t \in \mathbb{N}_0}$, where $(R_t)_{t \in [0, \infty)}$ is a discrete-time coalescent allowing for simultaneous multiple collisions of ancestral lineages.

Definition 2.1. *a) Let $R = (R_t)_{t \in [0, \infty)}$ be a continuous-time coalescent with simultaneous multiple collisions. We say that the considered population model is in the domain of attraction of R , if, for each sample size $n \in \mathbb{N}$, the time-scaled ancestral process $(\mathcal{R}_{[t/c_N]}^{(n)})_{t \in [0, \infty)}$ converges weakly to $(\varrho_n R_t)_{t \in [0, \infty)}$ as $N \rightarrow \infty$.*

b) Let $R = (R_t)_{t \in \mathbb{N}_0}$ be a discrete-time coalescent with simultaneous multiple collisions. We say that the considered population model is in the domain of attraction

of R , if, for each sample size $n \in \mathbb{N}$, the ancestral process $(\mathcal{R}_t^{(n)})_{t \in \mathbb{N}_0}$ converges weakly to $(\varrho_n R_t)_{t \in \mathbb{N}_0}$ as $N \rightarrow \infty$.

In both cases we call R the attractor of the considered population model.

Example 2.2. (extended Moran model) Let U be a random variable taking values in $\{0, \dots, N\}$. Define the offspring vector $\mu = (\mu_1, \dots, \mu_N)$ via $\mu_n := 1$ if $n \in \{1, \dots, N - U\}$, and, on $\{U > 0\}$, $\mu_n := 0$ if $n \in \{N - U + 1, \dots, N - 1\}$, and $\mu_N := U$. The associated shuffled exchangeable Cannings model was considered by Eldon and Wakeley [6]. For $U \equiv 0$ (and as well for $U \equiv 1$), we obtain the trivial model in which every individual has one offspring ($\mu_n = 1$ for all $n \in [N]$). For $U \equiv N$ in this model the N th individual is the parent of all the N children of the next generation ($\mu_N = N$). If $U \geq 1$ almost surely, then, by (2), the forward chain X has transition probabilities

$$\begin{aligned} P_{i,j} &= \frac{1}{\binom{N}{i}} \left(\sum_{1 \leq n_1 < \dots < n_i < N} \mathbb{P}(\sum_{k=1}^i \mu_{n_k} = j) + \sum_{1 \leq n_1 < \dots < n_i = N} \mathbb{P}(\sum_{k=1}^{i-1} \mu_{n_k} = j - U) \right) \\ &= \frac{1}{\binom{N}{i}} \mathbb{E} \left(\binom{N-U}{j} \binom{U-1}{i-j} + \binom{N-U}{j-U} \binom{U-1}{(i-1)-(j-U)} \right) \\ &= \begin{cases} \mathbb{E} \left(\binom{N-U}{j} \binom{U-1}{i-j} \right) / \binom{N}{i} & \text{if } j < i, \\ \mathbb{E} \left(\binom{N-U}{j} + \binom{N-U}{N-j} \right) / \binom{N}{i} = \mathbb{E} \left(\frac{(i)U + (N-i)U}{(N)U} \right) & \text{if } j = i, \\ \mathbb{E} \left(\binom{N-U}{N-j} \binom{U-1}{j-i} \right) / \binom{N}{i} & \text{if } j > i. \end{cases} \end{aligned}$$

If $\mathbb{P}(U = 0) > 0$, then in the above expressions for $P_{i,j}$ each U has to be replaced by $\max(U, 1)$. For $U \equiv 2$ this model reduces to the standard Moran model with forward transition probabilities $P_{i,i-1} = i(N-i)/(N)_2$, $i \in \{1, \dots, N\}$, $P_{i,i+1} = i(N-i)/(N)_2$, $i \in \{0, \dots, N-1\}$, $P_{i,i} = 1 - 2i(N-i)/(N)_2$, $i \in \{0, \dots, N\}$, and $P_{i,j} = 0$ otherwise. From (4), it follows similarly that, for $i, j \in [N]$,

$$\hat{P}_{i,j} = \begin{cases} \mathbb{E} \left(\binom{N-U}{j-1} \binom{U}{i-j+1} \right) / \binom{N}{i} & \text{if } j < i, \\ \mathbb{E} \left(\binom{N-U}{i} + U \binom{N-U}{i-1} \right) / \binom{N}{i} & \text{if } j = i, \\ 0 & \text{if } j > i. \end{cases}$$

Note that $\hat{P}_{i,1} = \mathbb{E}((U)_i)/(N)_i$, $i \in \{2, \dots, N\}$. In particular, $c_N = \hat{P}_{2,1} = \mathbb{E}((U)_2)/(N)_2$, in agreement with [6, Eq. (2)], and $d_N = \hat{P}_{3,1} = \mathbb{E}((U)_3)/(N)_3$. The model is in the domain of attraction of the Kingman coalescent if and only if $\mathbb{E}((U)_3)/(N\mathbb{E}((U)_2)) \rightarrow 0$ as $N \rightarrow \infty$. If U is binomially distributed with parameters N and $p \in [0, 1]$, then $c_N = p^2$. Note that $p = p_N$ may depend on N , so this model can have a wide variety of effective population sizes $N_e = 1/p_N^2$. For instance, if $p_N = N^{-\alpha}$, $\alpha > 0$, then $N_e = N^{2\alpha}$ is sub-linear for $\alpha < 1/2$ and super-linear for $\alpha > 1/2$. If $p_N = \lambda^N$, $\lambda < 1$, then $N_e = \lambda^{-2N}$ grows exponentially. We will come back to this model in Proposition 3.1.

In the following section we will introduce a skewed conditional branching process model, which can be viewed as a particular Cannings model. We will later identify the attractor R of several concrete such population models.

3. SKEWED CONDITIONAL BRANCHING PROCESS MODELS

Let $\xi_1, \xi_2, \dots$ be independent non-negative integer valued random variables and let f_n denote the probability generating function (pgf) of ξ_n , $n \in \mathbb{N}$. For any pgf g and $l \in \mathbb{N}_0$ we use the standard notation g^l for the l th power of g ($g^0 = 1$) and the notation $g^{(l)}$ for the l th derivative of g . Moreover, $[x^l]g(x)$ denotes the coefficient in front of x^l in the Taylor expansion of g around zero. For $N \in \mathbb{N}$ assume that $\mathbb{P}(\xi_1 + \dots + \xi_N = N) > 0$ and let $\mu = (\mu_1, \dots, \mu_N)$ be a random vector with distribution

$$\mathbb{P}(\mu = k) := \frac{\mathbb{P}(\xi_1 = k_1) \cdots \mathbb{P}(\xi_N = k_N)}{\mathbb{P}(\xi_1 + \dots + \xi_N = N)} = \frac{[x^{k_1}]f_1(x) \cdots [x^{k_N}]f_N(x)}{[x^N](f_1 \cdots f_N)(x)},$$

$k = (k_1, \dots, k_N) \in \Delta(N)$. The distribution of μ is hence that of $(\xi_1, \dots, \xi_N)$ conditioned on the event that $\xi_1 + \dots + \xi_N = N$. Note that, for $n \in [N]$, the marginal variable μ_n has distribution

$$\begin{aligned} \mathbb{P}(\mu_n = k) &= \frac{\mathbb{P}(\xi_n = k) \mathbb{P}(\sum_{m \in [N] \setminus \{n\}} \xi_m = N - k)}{\mathbb{P}(\xi_1 + \dots + \xi_N = N)} \\ &= \frac{([x^k]f_n(x))([x^{N-k}](\prod_{m \in [N] \setminus \{n\}} f_m(x)))}{[x^N](f_1 \cdots f_N)(x)}, \quad k \in \{0, \dots, N\}. \end{aligned}$$

For each $n \in [N]$ one may interpret μ_n as the number of offspring of individual n in a population with non-overlapping generations of constant population size N . Note that μ has pgf

$$(9) \quad \mathbb{E}(z_1^{\mu_1} \cdots z_N^{\mu_N}) = \frac{[x^N](f_1(xz_1) \cdots f_N(xz_N))}{[x^N](f_1 \cdots f_N)(x)}, \quad |z_1|, \dots, |z_N| \leq 1.$$

Let $l_1, \dots, l_N \in \mathbb{N}_0$ and put $l := l_1 + \dots + l_N$. Applying the ‘derivative operator’ $(\partial^{l_1}/\partial^{l_1} z_1) \cdots (\partial^{l_N}/\partial^{l_N} z_N)$ to the left-hand side and the right-hand side of (9), and noting that it is allowed to interchange this derivative operator with the ‘coefficient operator’ $[x^N]$ in the numerator on the right-hand side of (9), it follows that

$$\mathbb{E}((\mu_1)_{l_1} z_1^{\mu_1 - l_1} \cdots (\mu_N)_{l_N} z_N^{\mu_N - l_N}) = \frac{[x^{N-l}](\prod_{i=1}^N f_i^{(l_i)}(xz_i))}{[x^N](f_1 \cdots f_N)(x)}.$$

Taking the limit $z_1 \nearrow 1, \dots, z_N \nearrow 1$ shows that μ has descending factorial moments

$$(10) \quad \mathbb{E}((\mu_1)_{l_1} \cdots (\mu_N)_{l_N}) = \frac{[x^{N-l}](\prod_{i=1}^N f_i^{(l_i)}(x))}{[x^N](f_1 \cdots f_N)(x)}, \quad l_1, \dots, l_N \in \mathbb{N}_0.$$

If the random variables $\xi_1, \xi_2, \dots$ are identically distributed, then the model reduces to the conditional branching process model first introduced by Moran and Watterson [26] and further investigated for example by Karlin and McGregor [13]. In this case, for each fixed $N \in \mathbb{N}$, the random variables $\mu_1, \dots, \mu_N$ are exchangeable. Note however, that in general it is not assumed here that the random variables $\xi_1, \xi_2, \dots$ are identically distributed. In this sense the model is a bit more general than the conditional branching process model of [26]. Particular classes and examples of the model, which to the best of our knowledge are new or only briefly mentioned in the literature, are introduced and analyzed in the following sections, among them the skewed Wright-Fisher model (Section 5), the skewed Dirichlet model (Section 6), and the skewed Kimura model (Section 7). One may think that, at least for fixed N , the model can be reduced to a simpler model by randomly permutating

the random variables $\xi_1, \dots, \xi_N$. However, this is not the case. If $\eta_1, \dots, \eta_N$ denotes a random permutation of $\xi_1, \dots, \xi_N$, then the random variables $\eta_1, \dots, \eta_N$ are identically distributed but in general not independent (not even uncorrelated) anymore. In this sense the model is indeed more general than the conditional branching process model of Karlin and McGregor.

On the other hand, for arbitrary but fixed $N \in \mathbb{N}$, the model is well known from the literature in the following sense. Fix $N \in \mathbb{N}$ and let $\nu = (\nu_1, \dots, \nu_N)$ be a random permutation of $\mu = (\mu_1, \dots, \mu_N)$. As explained in Section 2, the model can be interpreted as an exchangeable Cannings model with population size N and offspring vector ν . Note that, for $k = (k_1, \dots, k_N) \in \Delta(N)$,

$$\begin{aligned} \frac{\mathbb{P}(\eta_1 = k_1, \dots, \eta_N = k_N)}{\mathbb{P}(\eta_1 + \dots + \eta_N = N)} &= \frac{1}{N!} \sum_{\pi} \frac{\mathbb{P}(\xi_{\pi 1} = k_1, \dots, \xi_{\pi N} = k_N)}{\mathbb{P}(\xi_1 + \dots + \xi_N = N)} \\ &= \frac{1}{N!} \sum_{\pi} \mathbb{P}(\pi \mu = k) = \mathbb{P}(\nu = k), \end{aligned}$$

so permutating the random variables $\xi_1, \dots, \xi_N$ (and conditioning) leads to the same exchangeable Cannings model as permutating the random variables $\mu_1, \dots, \mu_N$. Essentially all results known for exchangeable Cannings models apply to our model, which simplifies our further analysis significantly. For example, as explained in Section 2, the transition matrix of the forward process has entries (2). Noting that, for any subset M of $[N]$,

$$\begin{aligned} \mathbb{P}\left(\sum_{m \in M} \mu_m = j\right) &= \frac{\mathbb{P}(\sum_{m \in M} \xi_m = j) \mathbb{P}(\sum_{m \in [N] \setminus M} \xi_m = N - j)}{\mathbb{P}(\xi_1 + \dots + \xi_N = N)} \\ &= \frac{([x^j](\prod_{m \in M} f_m)(x))([x^{N-j}](\prod_{m \in [N] \setminus M} f_m)(x))}{[x^N](f_1 \cdots f_N)(x)}, \end{aligned}$$

it follows that the forward transition matrix $P = (P_{i,j})_{i,j \in S}$ has entries

$$(11) \quad P_{i,j} = \frac{1}{\binom{N}{i}} \sum_M \frac{([x^j](\prod_{m \in M} f_m)(x))([x^{N-j}](\prod_{m \in [N] \setminus M} f_m)(x))}{[x^N](f_1 \cdots f_N)(x)}, \quad i, j \in S,$$

where the sum $\sum_M$ extends over all subsets M of $[N]$ satisfying $|M| = i$. We now turn to the backward chain. From (4) and (10), it follows that the backward transition probabilities are of the form

$$(12) \quad \begin{aligned} \hat{P}_{i,j} &= \frac{1}{\binom{N}{i}} \sum_{1 \leq n_1 < \dots < n_j \leq N} \\ &\sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{[x^{N-i}](f_{n_1}^{(l_1)}(x) \cdots f_{n_j}^{(l_j)}(x) \prod_{m \in [N] \setminus \{n_1, \dots, n_j\}} f_m(x))}{l_1! \cdots l_j! [x^N](f_1 \cdots f_N)(x)}, \quad i, j \in S, \end{aligned}$$

with the convention that $\hat{P}_{i,0} = \delta_{i0}$, $i \in S$. In particular,

$$\hat{P}_{i,1} = \Phi_1(i) = \frac{1}{\binom{N}{i}} \sum_{n=1}^N \frac{[x^{N-i}](f_n^{(i)}(x) \prod_{m \in [N] \setminus \{n\}} f_m(x))}{[x^N](f_1 \cdots f_N)(x)}, \quad i \in S,$$

and the coalescence probability is given by

$$c_N = \hat{P}_{2,1} = \frac{1}{(N)_2} \sum_{n=1}^N \frac{[x^{N-2}](f_n''(x) \prod_{m \in [N] \setminus \{n\}} f_m(x))}{[x^N](f_1 \cdots f_N)(x)}.$$

Moreover, the eigenvalues $\lambda_i := \hat{P}_{i,i} = \Phi_i(1, \dots, 1)$, $i \in S$, of the matrix $\hat{P}$ are given by $\lambda_0 = 1$ and

$$\lambda_i = \frac{1}{\binom{N}{i}} \sum_{1 \leq n_1 < \dots < n_i \leq N} \frac{[x^{N-i}](f_{n_1}'(x) \cdots f_{n_i}'(x) \prod_{m \in [N] \setminus \{n_1, \dots, n_i\}} f_m(x))}{[x^N](f_1 \cdots f_N)(x)}$$

for $i \in \{1, \dots, N\}$.

Remark. One may write $\hat{P}_{i,j}$ in the form

$$\begin{aligned} \hat{P}_{i,j} &= \frac{1}{\binom{N}{i}} \frac{1}{[x^N](f_1 \cdots f_N)(x)} \\ &\quad \sum_{1 \leq n_1 < \dots < n_j \leq N} [x^{N-i}](S_{f_{n_1}, \dots, f_{n_j}, x}(i, j) \prod_{m \in [N] \setminus \{n_1, \dots, n_j\}} f_m(x)), \end{aligned}$$

where (see, for example, [23, Eq. (18)])

$$S_{g_1, \dots, g_j, x}(i, j) := \frac{i!}{j!} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{g_1^{(l_1)}(x)}{l_1!} \cdots \frac{g_j^{(l_j)}(x)}{l_j!}$$

are some sort of generalized Stirling numbers of the second kind, with the convention that, for $j = 0$, $S_{g_1, \dots, g_j, x}(i, j) = \delta_{i0}$, $i \in \mathbb{N}_0$.

Remark. The class of the conditional branching process models leads to a wide variety of Cannings models. However (see the following proposition), not all Cannings models are conditional branching process models.

Proposition 3.1. *The extended Moran model (see Example 2.2) is not a conditional branching process model, provided that $\mathbb{P}(2 \leq U \leq N-1) = 1$.*

Proof. Suppose that the extended Moran model with population size N is a conditional branching process model. Then, there exist i.i.d. random variables $\xi_1, \xi_2, \dots$ such that $\mathbb{P}(\xi_1 + \dots + \xi_N = N) \neq 0$ and $\mathbb{P}(\nu = k) = p_{k_1} \cdots p_{k_N} / \mathbb{P}(\xi_1 + \dots + \xi_N = N)$ for $k = (k_1, \dots, k_N) \in \Delta(N)$, where $p_k := \mathbb{P}(\xi_1 = k)$, $k \in \mathbb{N}_0$. In particular,

$$\begin{aligned} 0 &\neq \mathbb{E}\left(\frac{1}{N \binom{N-1}{U-1}}\right) \\ &= \mathbb{P}(\nu_1 = \dots = \nu_{N-U} = 1, \nu_{N-U+1} = \dots = \nu_{N-1} = 0, \nu_N = U) \\ &= \frac{\mathbb{E}(p_0^{U-1} p_1^{N-U} p_U)}{\mathbb{P}(\xi_1 + \dots + \xi_N = N)}, \end{aligned}$$

and, hence, $p_1 \neq 0$, since $\mathbb{P}(N-U \geq 1) = 1$ by assumption. On the other hand,

$$0 = \mathbb{P}(\nu_1 = \dots = \nu_N = 1) = \frac{\mathbb{P}(\xi_1 = 1) \cdots \mathbb{P}(\xi_N = 1)}{\mathbb{P}(\xi_1 + \dots + \xi_N = N)} = \frac{p_1^N}{\mathbb{P}(\xi_1 + \dots + \xi_N = N)},$$

and, hence, $p_1 = 0$, an obvious contradiction. In particular, for $N \geq 3$, the standard Moran model ($U \equiv 2$) is not a conditional branching process model. $\square$

It seems to be hard to derive further exact or asymptotic results as $N \rightarrow \infty$ for the general skewed conditional branching process model. We therefore focus in the following sections on important subclasses.

4. THE SKEWED COMPOUND POISSON CLASS

Let ϕ be a given power series of the form $\phi(z) = \sum_{m=1}^{\infty} \phi_m z^m / m!$, $|z| < r$ with positive radius $r \in (0, \infty]$ of convergence and with non-negative coefficients $\phi_m \geq 0$, $m \in \mathbb{N}$. We also assume that $\phi_1 > 0$. Let furthermore $\theta_1, \theta_2, \dots \in (0, \infty)$ be given strictly positive real parameters. In this section it is assumed that, for each $n \in \mathbb{N}$, the random variable ξ_n of the skewed conditional branching process model (as described in the previous section) has pgf

$$(13) \quad f_n(x) = \mathbb{E}(x^{\xi_n}) = \exp\left(-\theta_n \phi(z) \left(1 - \frac{\phi(zx)}{\phi(z)}\right)\right), \quad |x| \leq 1.$$

In (13), z is viewed as a fixed parameter. However, for the following approach we also see z as a variable satisfying $|z| < r$. In order to state the following lemma we need to introduce, for $\theta > 0$, the Taylor expansion

$$(14) \quad \exp(\theta \phi(z)) = \sum_{k=0}^{\infty} \frac{\sigma_k(\theta)}{k!} z^k, \quad |z| < r,$$

of $\exp(\theta \phi(z))$, seen as a function of z . Note that the coefficients $\sigma_k(\theta)$ are strictly positive and they depend on the sequence $\phi := (\phi_n)_{n \in \mathbb{N}}$. More precisely, the coefficients $\sigma_k(\theta)$ satisfy the recursion $\sigma_0(\theta) = 1$ and

$$(15) \quad \sigma_{k+1}(\theta) = \theta \sum_{l=0}^k \binom{k}{l} \phi_{k-l+1} \sigma_l(\theta), \quad k \in \mathbb{N}_0,$$

i.e. $\sigma_1(\theta) = \theta \phi_1$, $\sigma_2(\theta) = \theta \phi_2 + \theta^2 \phi_1^2$, $\sigma_3(\theta) = \theta \phi_3 + 3\theta^2 \phi_1 \phi_2 + \theta^3 \phi_1^3$, and so on. Note that, for each fixed $k \in \mathbb{N}$, $\sigma_k(\theta) \sim (\theta \phi_1)^k$ as $\theta \rightarrow \infty$.

Proposition 4.1. *If, for each $n \in \mathbb{N}$, the random variable ξ_n has a pgf of the form (13), then the forward process X of the associated skewed conditional branching process model has transition probabilities*

$$(16) \quad P_{i,j} = \frac{\binom{N}{j}}{\binom{N}{i}} \sum_M \frac{\sigma_j(\sum_{m \in M} \theta_m) \sigma_{N-j}(\sum_{m \in [N] \setminus M} \theta_m)}{\sigma_N(\Theta_N)}, \quad i, j \in S,$$

where $\Theta_N := \theta_1 + \dots + \theta_N$, the sum $\sum_M$ extends over all subsets $M \subseteq [N]$ satisfying $|M| = i$, and the coefficients $\sigma_k(\theta)$ are recursively defined via (15).

Proof. For $j \in [N]$, $\theta > 0$ and $|x| \leq 1$, it follows from (14) that

$$[x^j] \exp(\theta \phi(zx)) = [x^j] \sum_{k=0}^{\infty} \frac{\sigma_k(\theta)}{k!} (zx)^k = z^j \frac{\sigma_j(\theta)}{j!}.$$

Using, for $M \subseteq [N]$, the shortage $\theta := \sum_{m \in M} \theta_m$, it follows that

$$\begin{aligned} [x^j] \left(\sum_{m \in M} f_m \right)(x) &= [x^j] \exp\left(-\theta \phi(z) \left(1 - \frac{\phi(zx)}{\phi(z)}\right)\right) \\ &= \exp(-\theta \phi(z)) [x^j] \exp(\theta \phi(zx)) = \exp(-\theta \phi(z)) z^j \frac{\sigma_j(\theta)}{j!}. \end{aligned}$$

Thus, (16) follows from (11). $\square$

Remark. For the unbiased case, when the parameter $\theta_n = \theta$ does not depend on $n \in \mathbb{N}$, (16) reduces to

$$P_{i,j} = \binom{N}{j} \frac{\sigma_j(i\theta) \sigma_{N-j}((N-i)\theta)}{\sigma_N(N\theta)}, \quad i, j \in S.$$

Let us now turn to the backward process.

Proposition 4.2. *If, for each $n \in \mathbb{N}$, the random variable ξ_n has a pgf of the form (13), then the backward process $\hat{X}$ of the associated skewed conditional branching process model has transition probabilities*

$$(17) \quad \hat{P}_{i,j} = \frac{i!}{\sigma_i(\Theta_N)} \sum_{1 \leq n_1 < \dots < n_j \leq N} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{\sigma_{l_1}(\theta_{n_1}) \cdots \sigma_{l_j}(\theta_{n_j})}{l_1! \cdots l_j!}, \quad i, j \in S,$$

with the convention that $\hat{P}_{i,0} = \delta_{i,0}$, $i \in S$. Here $\Theta_N := \theta_1 + \dots + \theta_N$ and the coefficients $\sigma_k(\theta)$ are recursively defined via (15). In particular,

$$\hat{P}_{i,1} = \frac{1}{\sigma_i(\Theta_N)} \sum_{n=1}^N \sigma_i(\theta_n), \quad i \in \{1, \dots, N\}.$$

Proof. From Kolchin's representation formula [15] (see also [28, Chapter 1, Theorem 1.2]), it follows that $\mu = (\mu_1, \dots, \mu_N)$ has distribution

$$\mathbb{P}(\mu = k) = \frac{N!}{\sigma_N(\Theta_N)} \prod_{n=1}^N \frac{\sigma_{k_n}(\theta_n)}{k_n!}, \quad k = (k_1, \dots, k_N) \in \Delta(N).$$

Therefore, μ has joint descending factorial moments

$$\mathbb{E}((\mu_1)_{l_1} \cdots (\mu_N)_{l_N}) = \frac{(N)_{l_1 + \dots + l_N}}{\sigma_{l_1 + \dots + l_N}(\Theta_N)} \prod_{n=1}^N \sigma_{l_n}(\theta_n), \quad l_1, \dots, l_N \in \mathbb{N}_0.$$

The probability (6) is therefore of the form

$$\begin{aligned} \Phi_j(l_1, \dots, l_j) &= \frac{1}{(N)_{l_1 + \dots + l_j}} \sum_{\substack{n_1, \dots, n_j = 1 \\ \text{all distinct}}}^N \mathbb{E}((\mu_{n_1})_{l_1} \cdots (\mu_{n_j})_{l_j}) \\ &= \frac{1}{\sigma_{l_1 + \dots + l_j}(\Theta_N)} \sum_{\substack{n_1, \dots, n_j = 1 \\ \text{all distinct}}}^N \sigma_{l_1}(\theta_{n_1}) \cdots \sigma_{l_j}(\theta_{n_j}), \quad l_1, \dots, l_j \in \mathbb{N}. \end{aligned}$$

Using (7), it follows that

$$\begin{aligned}
\hat{P}_{i,j} &= \frac{i!}{j!} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{\Phi_j(l_1, \dots, l_j)}{l_1! \dots l_j!} \\
&= \frac{i!}{j! \sigma_i(\Theta_N)} \sum_{\substack{n_1, \dots, n_j = 1 \\ \text{all distinct}}}^N \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{\sigma_{l_1}(\theta_{n_1}) \dots \sigma_{l_j}(\theta_{n_j})}{l_1! \dots l_j!} \\
&= \frac{i!}{\sigma_i(\Theta_N)} \sum_{1 \leq n_1 < \dots < n_j \leq N} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{\sigma_{l_1}(\theta_{n_1}) \dots \sigma_{l_j}(\theta_{n_j})}{l_1! \dots l_j!}. \quad \square
\end{aligned}$$

Remark. The previous proposition in particular shows that

$$c_N = \hat{P}_{2,1} = \frac{1}{\sigma_2(\Theta_N)} \sum_{n=1}^N \sigma_2(\theta_n) = \frac{\phi_2 \Theta_N + \phi_1^2 \sum_{n=1}^N \theta_n^2}{\phi_2 \Theta_N + \phi_1^2 \Theta_N^2}.$$

Since $N \sum_{n=1}^N \theta_n^2 \geq \Theta_N^2$, it follows that $c_N \geq 1/N$, or, equivalently, that $N_e \leq N$.

Let us now focus on the unbiased case, when all the parameters $\theta_n = \theta$ are equal to some constant $\theta \in (0, \infty)$. It turns out to be convenient to introduce, for any formal series $a(z) = \sum_{k=1}^{\infty} a_k z^k / k!$ or, equivalently, for any sequence $a. := (a_1, a_2, \dots)$, the Bell polynomials

$$(18) \quad B_{i,j}(a.) := \frac{i!}{j!} [z^i] (a(z))^j = \frac{i!}{j!} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{a_{l_1} \dots a_{l_j}}{l_1! \dots l_j!}, \quad i, j \in \mathbb{N}_0.$$

Note that $B_{0,0}(a.) = 1$, $B_{i,0}(a.) = B_{0,i}(a.) = 0$, $B_{i,1}(a.) = a_i$ and $B_{i,i}(a.) = a_1^i$ for $i \in \mathbb{N}$. For more information on these polynomials we refer the reader to [1] and Chapter 1 of Pitman [28].

The following theorem provides exact and asymptotic formulae for the transition probabilities $\hat{P}_{i,j}$ and clarifies that the unbiased compound Poisson class is in the domain of attraction of the Kingman coalescent.

Theorem 4.3. *If $\theta_n = \theta \in (0, \infty)$ for all $n \in \mathbb{N}$, then*

$$(19) \quad \hat{P}_{i,j} = \frac{(N)_j}{\sigma_i(\theta N)} B_{i,j}(\sigma.(\theta)) \quad i, j \in S,$$

where the $B_{i,j}(\sigma.(\theta))$ are the Bell polynomials of $\sigma.(\theta) := (\sigma_1(\theta), \sigma_2(\theta), \dots)$. In particular,

$$(20) \quad \hat{P}_{i,1} = \frac{N \sigma_i(\theta)}{\sigma_i(\theta N)}, \quad i \in \{1, \dots, N\}.$$

Moreover, for $i, j \in \mathbb{N}$ with $i \geq j$,

$$(21) \quad \hat{P}_{i,j} \sim \frac{B_{i,j}(\sigma.(\theta))}{(\theta \phi_1)^i} \frac{1}{N^{i-j}}, \quad N \rightarrow \infty,$$

and the model is in the domain of attraction of the Kingman coalescent in the sense of Definition 2.1 a).

Proof. Since $\theta_n = \theta$ does not depend on n , the formula (17) reduces to (19) thanks to the formula (18) for the Bell polynomials. For $j = 1$, (19) reduces to (20), since $B_{i,1}(\sigma(\theta)) = \sigma_i(\theta)$. For $i, j \in \mathbb{N}$ with $i \geq j$ and for all $N \geq i$ it follows that

$$\hat{P}_{i,j} = \frac{(N)_j}{\sigma_i(\theta N)} B_{i,j}(\sigma(\theta)) \sim \frac{N^j}{(\theta N \phi_1)^i} B_{i,j}(\sigma(\theta)) = \frac{B_{i,j}(\sigma(\theta))}{(\theta \phi_1)^i} \frac{1}{N^{i-j}},$$

which is (21). For $j = 1$, Eq. (21) reduces to $\hat{P}_{i,1} \sim \sigma_i(\theta)/((\theta \phi_1)^i N^{i-1})$ as $N \rightarrow \infty$. Thus, $d_N/c_N = \hat{P}_{3,1}/\hat{P}_{2,1} \sim (\sigma_3(\theta)/(\theta \phi_1 \sigma_2(\theta)) N^{-1} \rightarrow 0$, which ensures that the model is in the domain of attraction of the Kingman coalescent. $\square$

Remark. Theorem 4.3 in particular provides explicit exact formulae for the coalescence probability $c_N = \hat{P}_{2,1}$ and for $d_N = \hat{P}_{3,1}$, namely

$$(22) \quad c_N = \frac{N \sigma_2(\theta)}{\sigma_2(\theta N)} = \frac{N(\theta \phi_2 + \theta^2 \phi_1^2)}{\theta N \phi_2 + (\theta N)^2 \phi_1^2}$$

and

$$d_N = \frac{N \sigma_3(\theta)}{\sigma_3(\theta N)} = \frac{N(\theta \phi_3 + 3\theta^2 \phi_1 \phi_2 + \theta^3 \phi_1^3)}{\theta N \phi_3 + 3(\theta N)^2 \phi_1 \phi_2 + (\theta N)^3 \phi_1^3}.$$

Example 4.4. (Wright-Fisher model) For the Wright-Fisher model, $\phi(z) = z$, $\sigma_k(\theta) = \theta^k$. From $B_{i,j}(\theta, \theta^2, \dots) = \theta^i B_{i,j}(1, 1, \dots) = \theta^i S(i, j)$, where the $S(i, j)$ are the Stirling numbers of the second kind, it follows that

$$\hat{P}_{i,j} = \frac{(N)_j}{\sigma_i(\theta N)} B_{i,j}(\sigma(\theta)) = \frac{(N)_j}{(\theta N)^i} \theta^i S(i, j) = (N)_j N^{-i} S(i, j),$$

which is well known. For results concerning the skewed Wright-Fisher model we refer the reader to Section 5.

Example 4.5. (Dirichlet model) Suppose that $\phi(z) = -\log(1 - z)$, $|z| < 1$, i.e. $\phi_m = (m - 1)!$, $m \in \mathbb{N}$. Then, $\sigma_k(\theta) = [\theta]_k$, $k \in \mathbb{N}$, and, hence,

$$\hat{P}_{i,j} = \frac{(N)_j}{[\theta N]_i} B_{i,j}([\theta]_\cdot) = \frac{i!}{[\theta N]_i} \binom{N}{j} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{[\theta]_{l_1} \cdots [\theta]_{l_j}}{l_1! \cdots l_j!}.$$

Results on the skewed Dirichlet model are presented in Section 6.

We briefly mention two further examples, showing the wide variety of models we are concerned with.

Example 4.6. Let $\alpha \in (0, 1]$ and assume that $\phi(z) = 1 - (1 - z)^\alpha$, $|z| < 1$. Then, $\phi_m = (-1)^{m-1}(\alpha)_m = \alpha[1 - \alpha]_{m-1} \geq 0$, $m \in \mathbb{N}$. Note that $\phi(z) \rightarrow 1$ as $z \rightarrow 1$, a smoothness property of ϕ . For $z \rightarrow 1$ the random variable ξ_1 has a discrete stable distribution with pgf $\mathbb{E}(x^{\xi_1}) = \exp(-\theta(1 - x)^\alpha)$ (see, for example, Steutel and van Harn [33, Eq. (3.7)]) and tail asymptotics $\mathbb{P}(\xi_1 > y) \sim \theta c_\alpha y^{-\alpha}$ as $y \rightarrow \infty$, where $c_\alpha := \sin(\pi\alpha/2) \cos(\pi\alpha/2) \Gamma(\alpha)/(\pi/2)$ (apply, for example, Pitman [27, Theorem 1]). Note that $\phi_1 = \alpha$ and that $\phi_2 = \alpha\beta$ with $\beta := 1 - \alpha$. Thus, from (22), it follows that

$$c_N = \hat{P}_{2,1} = \frac{N(\theta\alpha\beta + \theta^2\alpha^2)}{\theta N\alpha\beta + (\theta N)^2\alpha^2} \sim \left(1 + \frac{\beta}{\theta\alpha}\right) \frac{1}{N}.$$

Thus, $N_e = 1/c_N \sim \rho N$ with $\rho := (1 + \beta/(\theta\alpha))^{-1} < 1$.

Example 4.7. Let $\alpha \in (0, \infty)$ and assume that $\phi(z) = (1 - z)^{-\alpha} - 1$, $|z| < 1$. Then, $\phi_m = [\alpha]_m$, $m \in \mathbb{N}$. Note that $\phi(z) \rightarrow \infty$ as $z \rightarrow 1$. The random variable ξ_1 has a compound Poisson distribution with negative binomial jumps and pgf $\mathbb{E}(x^{\xi_1}) = \exp(-\theta\phi(z)(1 - \phi(zx)/\phi(z)))$, $|x| \leq 1$. By (22),

$$c_N = \frac{N(\theta\alpha(\alpha + 1) + \theta^2\alpha^2)}{\theta N\alpha(\alpha + 1) + (\theta N)^2\alpha^2} \sim \left(1 + \frac{\alpha + 1}{\theta\alpha}\right) \frac{1}{N}.$$

Thus, $N_e \sim \rho N$ with $\rho := (1 + (1 + \alpha)/(\theta\alpha))^{-1} < 1$.

Theorem 4.3 clarifies that a large class of unbiased conditional branching process models is in the domain of attraction of the Kingman coalescent. For the skewed situation the asymptotical behavior of (17) as $N \rightarrow \infty$ is much more involved. In the following sections we focus on particular skewed population models. It will turn out that these models are not necessarily in the domain of attraction of the Kingman coalescent.

5. SKEWED WRIGHT-FISHER MODEL

Let $m_1, m_2, \dots \in (0, \infty)$ be given parameters and assume that ξ_n is Poisson distributed with parameter m_n , $n \in \mathbb{N}$. Since ξ_n has pgf $f_n(x) = \mathbb{E}(x^{\xi_n}) = e^{-m_n(1-x)}$, $n \in \mathbb{N}$, it follows that this model belongs to the skewed compound Poisson class (13) with $\phi(z) := z$ and $\theta_n := m_n/z$. Moreover, from (9) and (10) it follows that μ has a multinomial distribution with parameters N and $s_{1,N}, \dots, s_{N,N}$, where $s_{n,N} := m_n/(m_1 + \dots + m_N)$, $n \in \{1, \dots, N\}$. We shall often drop the index N and use s_n instead of $s_{n,N}$ for notational convenience. Note that μ has pgf $\mathbb{E}(z_1^{\mu_1} \dots z_N^{\mu_N}) = (s_1 z_1 + \dots + s_N z_N)^N$, $|z_1|, \dots, |z_N| \leq 1$, and descending factorial moments $\mathbb{E}((\mu_1)_{l_1} \dots (\mu_N)_{l_N}) = (N)_l s_1^{l_1} \dots s_N^{l_N}$, $l_1, \dots, l_N \in \mathbb{N}_0$, where $l := l_1 + \dots + l_N$. The pgf of the n th marginal μ_n is $\mathbb{E}(z_n^{\mu_n}) = (1 - s_n + z_n s_n)^N$, $|z_n| \leq 1$, showing that μ_n has a binomial distribution with parameters N and s_n . In particular, $\mathbb{E}(\mu_n) = N s_n$ and $\sigma_n^2 := \text{Var}(\mu_n) = N s_n(1 - s_n)$, $n \in \{1, \dots, N\}$. From the above formula for the descending factorial moments of μ , it follows that $\mathbb{E}(\mu_{n_1} \mu_{n_2}) = N(N-1) s_{n_1} s_{n_2}$ for $n_1, n_2 \in \{1, \dots, N\}$ with $n_1 \neq n_2$. In particular, $\text{Cov}(\mu_{n_1}, \mu_{n_2}) = -N s_{n_1} s_{n_2}$ for $n_1 \neq n_2$.

The expression (11) for the forward transition probabilities $P_{i,j}$ simplifies to

$$(23) \quad P_{i,j} = \frac{1}{\binom{N}{i}} \sum_{1 \leq n_1 < \dots < n_i \leq N} \binom{N}{j} \left(\sum_{k=1}^i s_{n_k} \right)^j \left(1 - \sum_{k=1}^i s_{n_k} \right)^{N-j}, \quad i, j \in S.$$

Similarly, (12) reduces to

$$(24) \quad \hat{P}_{i,j} = i! \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{1}{l_1! \dots l_j!} \sum_{1 \leq n_1 < \dots < n_j \leq N} s_{n_1}^{l_1} \dots s_{n_j}^{l_j}, \quad i, j \in S,$$

which, for $j = 1$, yields $\hat{P}_{i,1} = \sum_{n=1}^N s_n^i$, $i \in \{1, \dots, N\}$. For $i = 2$ and $i = 3$, we obtain $c_N = \hat{P}_{2,1} = \sum_{n=1}^N s_n^2$ and $d_N = \hat{P}_{3,1} = \sum_{n=1}^N s_n^3$. Note that $c_N = \sum_{n=1}^N s_n^2 \geq 1/N$. In particular, $N_e := 1/c_N \leq N$, so the effective population size is smaller than or equal to the effective population size in the unbiased case. Applying the Hölder inequality $\sum_n |a_n b_n| \leq (\sum_n a_n^2)^{1/2} (\sum_n b_n^2)^{1/2}$ with $a_n := s_n^{1/2}$ and

$b_n := s_n^{3/2}$ shows that $c_N \leq d_N^{1/2}$, or, equivalently, that $c_N \leq d_N/c_N$. On the other hand, since the Euclidian 2-norm is larger than or equal to the Euclidian 3-norm, we obtain $c_N^{1/2} \geq d_N^{1/3}$, or, equivalently, $d_N/c_N \leq c_N^{1/2}$. Thus, $c_N \leq d_N/c_N \leq c_N^{1/2}$ and, consequently, $c_N \rightarrow 0$ is equivalent to $d_N/c_N \rightarrow 0$. The model is hence in the domain of attraction of the Kingman coalescent if and only if $c_N \rightarrow 0$.

Choosing $i = j$ in (24) shows that the backward matrix $\hat{P}$ has eigenvalues $\lambda_0 = 1$ and

$$\lambda_i := \hat{P}_{i,i} = i! \sum_{1 \leq n_1 < \dots < n_i \leq N} s_{n_1} \cdots s_{n_i}, \quad i \in \{1, \dots, N\}$$

By duality [20], these are also the eigenvalues of the forward transition matrix P . Note that $1 = \lambda_0 = \lambda_1 > \lambda_2 = 1 - c_N > \lambda_3 > \dots > \lambda_N$. In particular, the eigenvalues $\lambda_1, \dots, \lambda_N$ are pairwise distinct. Coming back to the random variables $\mu_1, \dots, \mu_N$ we see that the average covariances of the μ_n 's is

$$\frac{1}{\binom{N}{2}} \sum_{1 \leq n_1 < n_2 \leq N} \text{Cov}(\mu_{n_1}, \mu_{n_2}) = -\frac{2}{N-1} \sum_{1 \leq n_1 < n_2 \leq N} s_{n_1} s_{n_2} = -\frac{1 - c_N}{N-1}.$$

Let

$$S_N := \frac{1}{N} \sum_{n=1}^N \frac{\mathbb{E}(\mu_n)}{N} \mu_n = \frac{1}{N} \sum_{n=1}^N s_n \mu_n$$

be the size-biased relative mean offspring fertility. Since μ has a multinomial distribution with parameters N and $s_1, \dots, s_N$, it is readily checked that $\mathbb{E}(S_N) = c_N$ and that $\text{Var}(S_N) = (d_N - c_N^2)/N$, which provides an alternative proof of the inequality $d_N \geq c_N^2$.

For the Wright-Fisher model, the moments of X_{t+1} , given $X_t = i$, can be related to the backward probabilities as follows. For $j \in \mathbb{N}$, we have

$$\begin{aligned} \mathbb{E}(X_{t+1}^j \mid X_t = i) &= \frac{1}{\binom{N}{i}} \sum_{1 \leq n_1 < \dots < n_i \leq N} \mathbb{E}\left(\left(\sum_{k=1}^i \mu_{n_k}\right)^j\right) \\ &= \frac{1}{\binom{N}{i}} \sum_{l=1}^j S(l, j)(N)_l \sum_{1 \leq n_1 < \dots < n_i \leq N} \left(\sum_{k=1}^i s_{n_k}\right)^l. \end{aligned}$$

Using

$$\sum_{1 \leq n_1 < \dots < n_i \leq N} \left(\sum_{k=1}^i s_{n_k}\right)^l = \sum_{k=1}^l \binom{N-k}{i-k} \hat{P}_{l,k},$$

where $\hat{P}_{l,k}$ is defined in (24), we get

$$\begin{aligned} \mathbb{E}(X_{t+1}^j \mid X_t = i) &= \frac{1}{\binom{N}{i}} \sum_{l=1}^j S(l, j)(N)_l \sum_{k=1}^l \binom{N-k}{i-k} \hat{P}_{l,k} \\ &= \sum_{l=1}^j S(l, j)(N)_l \sum_{k=1}^l \frac{(i)_k}{(N)_k} \hat{P}_{l,k}, \end{aligned}$$

which is a polynomial of degree j in i .

For the Wright-Fisher model, there are the following two alternative representations of the backward transition probabilities (24).

(i) It can easily be checked that

$$(25) \quad \hat{P}_{i,j} = N^{-i}(N)_j S(i, j, Ns),$$

where $s := (s_1, \dots, s_N)$ and

$$(26) \quad S(i, j, s) := \frac{i!}{j!} [x^i] \frac{1}{\binom{N}{j}} \sum_{1 \leq n_1 < \dots < n_j \leq N} \prod_{k=1}^j (e^{xs_{n_k}} - 1)$$

are generalized second kind Stirling numbers. Note that, when $\mathbb{E}(\mu_n) = Ns_n = 1$ for all $n \in [N]$, (26) reduces to the usual second kind Stirling numbers $S(i, j) = (i!/j!)[x^i](e^x - 1)^j$ (see [5, vol. I, p. 144]). The unbiased version of (25) is thus $\hat{P}_{i,j} = N^{-i}(N)_j S(i, j)$, which is well-known. Note that (25) and (26) also yield

$$\hat{P}_{i,j} = i! [x^i z^j] \prod_{n=1}^N (1 + z(e^{xs_n} - 1))$$

showing that, as shown in [18],

$$\mathbb{E}(z^{\hat{X}_{t+1}} | \hat{X}_t = i) = i! [x^i] \prod_{n=1}^N (1 + z(e^{xs_n} - 1)),$$

and for the double pgf

$$\sum_i \frac{x^i}{i!} \mathbb{E}(z^{\hat{X}_{t+1}} | \hat{X}_t = i) = \prod_{n=1}^N (1 + z(e^{xs_n} - 1)).$$

This leads in particular to $\mathbb{E}(\hat{X}_{t+1} | \hat{X}_t = i) = \sum_{n=1}^N (1 - (1 - s_n)^i)$ and

$$\begin{aligned} & \mathbb{E}\left(\binom{\hat{X}_{t+1}}{2} \mid \hat{X}_t = i\right) \\ &= \sum_{n=1}^N (n-1)(1 - (1 - s_n)^i) - \sum_{1 \leq n_2 < n_1 \leq N} ((1 - s_{n_1})^i - (1 - (s_{n_1} + s_{n_2}))^i). \end{aligned}$$

(ii) There is a second obvious representation taking into account repetitions, in the spirit of the Ewens sampling formula [8]. Assume there are a_l individuals at generation t , numbered $1 \leq n_{1,l} < \dots < n_{a_l,l} \leq N$, producing exactly l offspring, $l \in \{0, \dots, N\}$. Clearly, there is no overlap of the above number of sequences for different values of l . Then, using (24)

$$(27) \quad \hat{P}_{i,j} = i! \sum_{\substack{\sum_{l=1}^i l a_l = i; \\ \sum_{l=1}^i a_l = j}} \sum_{l=1}^i \sum_{n_{1,l} < \dots < n_{a_l,l} \leq N} \prod_{l=1}^i \prod_{k=1}^{a_l} \left(\frac{s_{n_{k,l}}^l}{l!} \right).$$

In (27), there are

$$\prod_{l=1}^i \binom{N - \sum_{k=1}^{l-1} a_k}{a_l} = \frac{(N)_j}{\prod_{l=1}^i a_l!}$$

sums of the type $\sum_{l=1}^i \sum_{n_{1,l} < \dots < n_{a_l,l}} 1$ so that when $s_n = 1/N$ for all $n \in [N]$ (the unbiased case),

$$\hat{P}_{i,j} = N^{-i}(N)_j \sum_{\substack{\sum_{l=1}^i la_l = i; \\ \sum_{l=1}^i a_l = j}} \frac{i!}{\prod_{l=1}^i a_l! l^{a_l}},$$

where the last sum is an alternative representation of the second kind Stirling numbers $S(i, j)$ (see [5, vol. I, p. 145]).

Example 5.1. (power law growth) Fix a constant $\alpha \in \mathbb{R}$ and assume that $m_n := \mathbb{E}(\xi_n) = n^{-\alpha}$, $n \in \mathbb{N}$. The unbiased case ($m_n = 1$ for all $n \in \mathbb{N}$) corresponds to $\alpha = 0$. In the following seven ranges for the parameter α are distinguished.

(i) If $\alpha < 1/3$, then $M_N := \sum_{n=1}^N m_n \sim N \int_0^1 (Nx)^{-\alpha} dx = N^{1-\alpha}/(1-\alpha)$,

$$(28) \quad c_N = \frac{1}{M_N^2} \sum_{n=1}^N n^{-2\alpha} \sim \left(\frac{1-\alpha}{N^{1-\alpha}} \right)^2 \frac{N^{1-2\alpha}}{1-2\alpha} = \frac{(1-\alpha)^2}{(1-2\alpha)N}$$

and, similarly,

$$d_N = \frac{1}{M_N^3} \sum_{n=1}^N n^{-3\alpha} \sim \left(\frac{1-\alpha}{N^{1-\alpha}} \right)^3 \frac{N^{1-3\alpha}}{1-3\alpha} = \frac{(1-\alpha)^3}{1-3\alpha} \frac{1}{N^2}.$$

(ii) If $\alpha = 1/3$, then (28) still holds, i.e. $c_N \sim 4/(3N)$, but

$$d_N = \frac{1}{M_N^3} \sum_{n=1}^N \frac{1}{n} \sim \left(\frac{1-\alpha}{N^{1-\alpha}} \right)^3 \log N = \frac{8}{27} \frac{\log N}{N^2}.$$

(iii) If $1/3 < \alpha < 1/2$, then (28) still holds, but

$$d_N = \frac{1}{M_N^3} \sum_{n=1}^N n^{-3\alpha} \sim \left(\frac{1-\alpha}{N^{1-\alpha}} \right)^3 \zeta(3\alpha) = (1-\alpha)^3 \zeta(3\alpha) \frac{1}{N^{3(1-\alpha)}},$$

where ζ denotes the Riemann zeta function.

In all three cases, i.e. for $\alpha < 1/2$, we have $N_e = 1/c_N \sim \rho N$ with $0 < \rho := (1-2\alpha)/(1-\alpha)^2 \leq 1$. The effective population size is hence asymptotically of a factor ρ smaller than the effective population size N in the unbiased case ($\alpha = 0$).

Moreover, from (24), it follows that the eigenvalues satisfy

$$\lambda_i = \hat{P}_{i,i} = \frac{i!}{M_N^i} \sum_{1 \leq n_1 < \dots < n_i \leq N} (n_1 \dots n_i)^{-\alpha} \sim \left(\frac{1-\alpha}{N^{1-\alpha}} \right)^i i! \sum_{n=i!}^{(N)_i} n^{-\alpha} M_N(n, i),$$

where $M_N(n, i)$ is the number of multiplicative partitions of n into i ordered distinct factors each belonging to $\{1, \dots, N\}$.

We will now see that N_e can increase of order slower than N .

(iv) When $\alpha = 1/2$, it is readily checked that $M_N \sim 2N^{1/2}$, $c_N \sim (\log N)/(4N)$, and $d_N \sim \zeta(3/2)/(8N^{3/2})$. Thus, $N_e = 1/c_N \sim (4N)/\log N$ is asymptotically of a factor $4/\log N$ smaller than the standard effective population size N in the unbiased case.

(v) Assume now that $1/2 < \alpha < 1$. Then, still $M_N = \sum_{n=1}^N n^{-\alpha} \sim N^{1-\alpha}/(1-\alpha)$,

$$c_N = \frac{1}{M_N^2} \sum_{n=1}^N n^{-2\alpha} \sim \frac{(1-\alpha)^2}{N^{2(1-\alpha)}} \zeta(2\alpha)$$

and, similarly,

$$d_N = \frac{1}{M_N^3} \sum_{n=1}^N n^{-3\alpha} \sim \frac{(1-\alpha)^3}{N^{3(1-\alpha)}} \zeta(3\alpha)$$

Note that, with $\rho := 1/((1-\alpha)^2 \zeta(2\alpha)) < 1$, $N_e \sim \rho N^{2(1-\alpha)}$ grows algebraically and the order is slower than N .

(vi) For $\alpha = 1$ it is straightforward to check that $M_N \sim \log N$, $c_N \sim \zeta(2)/(\log N)^2$, and $d_N \sim \zeta(3)/(\log N)^3$. Note that $N_e \sim (\log N)^2/\zeta(2)$ grows quite slow (logarithmically).

For all six cases (i) - (vi) considered so far, i.e. for $\alpha \leq 1$, we have $c_N \rightarrow 0$ and $d_N/c_N \rightarrow 0$. Thus, in the sense of Definition 2.1 a), the model is in the domain of attraction of the Kingman coalescent. Thus, for each sample size $n \in \mathbb{N}$, the time-scaled backward process $(\mathcal{R}_{[t/c_N]}^{(n)})_{t \in [0, \infty)}$ converges weakly to the Kingman n -coalescent as $N \rightarrow \infty$. All that is left about the details of the original discrete fertility model ($m_n = n^{-\alpha}$) is enclosed in the parameter $N_e = 1/c_N \leq N$, where N is the time-scale of the unbiased Wright-Fisher model ($m_n = 1$).

(vii) Assume now that $\alpha > 1$. Then, $M_N = \sum_{n=1}^N n^{-\alpha} \rightarrow \zeta(\alpha) =: M$, and,

$$\hat{P}_{i,1} = \Phi_1(i) = \sum_{n=1}^N s_n^i = \frac{1}{M_N^i} \sum_{n=1}^N n^{-i\alpha} \rightarrow \frac{\zeta(i\alpha)}{(\zeta(\alpha))^i}, \quad i \in \mathbb{N}.$$

In particular $c_N \rightarrow \zeta(2\alpha)/(\zeta(\alpha))^2 > 0$ and $d_N \rightarrow \zeta(3\alpha)/(\zeta(\alpha))^3 > 0$. The regime $\alpha > 1$ thus differs significantly from the previously studied cases, since c_N converges to a positive constant as $N \rightarrow \infty$. More generally, for $l_1, \dots, l_j \in \mathbb{N}$, we have to analyze the behavior of

$$\Phi_j^{(N)}(l_1, \dots, l_j) = \sum_{\substack{n_1, \dots, n_j=1 \\ \text{all distinct}}}^N s_{n_1, N}^{l_1} \cdots s_{n_j, N}^{l_j}$$

as $N \rightarrow \infty$. For each fixed $n \in \mathbb{N}$ we have $s_{n, N} = n^{-\alpha}/M_N \rightarrow n^{-\alpha}/M = n^{-\alpha}/\zeta(\alpha) =: p_n$ as $N \rightarrow \infty$. Note that $\sum_{n=1}^{\infty} p_n = 1$. Moreover, $s_{n, N} \leq n^{-\alpha}$ uniformly for all N , since $M_N \geq 1$. Thus, by dominated convergence, for each $l_1, \dots, l_j \in \mathbb{N}$, the limit $\phi_j(l_1, \dots, l_j) := \lim_{N \rightarrow \infty} \Phi_j^{(N)}(l_1, \dots, l_j)$ exists and is of the form

$$(29) \quad \phi_j(l_1, \dots, l_j) = \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} p_{n_1}^{l_1} \cdots p_{n_j}^{l_j}.$$

For arbitrary $i, j \in \mathbb{N}$ it therefore follows from (24) that

$$(30) \quad \lim_{N \rightarrow \infty} \hat{P}_{i,j} = \frac{i!}{j!} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \frac{\phi_j(l_1, \dots, l_j)}{l_1! \cdots l_j!}.$$

It is convenient (see, Schweinsberg [31]) to rewrite (29) in integral form as

$$\phi_j(l_1, \dots, l_j) = \int_{\Delta} \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} x_{n_1}^{l_1} \cdots x_{n_j}^{l_j} \frac{\Xi(dx)}{(x, x)},$$

where $\Delta := \{(x_1, x_2, \dots) : x_1 \geq x_2 \geq \dots \geq 0, \sum_{n=1}^{\infty} x_n \leq 1\}$, $(x, x) := \sum_{n=1}^{\infty} x_n^2$ for $x \in \Delta$ and the measure Ξ on Δ assigns its total mass $\Xi(\Delta) := (p, p) = \sum_{n=1}^{\infty} p_n^2 = \zeta(2\alpha)/(\zeta(\alpha))^2$ to the single point $p = (p_1, p_2, \dots) \in \Delta$. Theorem 2.1 of [24] ensures that the model, without any time-scaling involved, is in the domain of attraction (in the sense of Definition 2.1 b)) of the discrete-time Ξ -coalescent with the measure Ξ as just defined.

Example 5.2. (geometric growth) Fix a constant $\lambda \in (0, \infty)$ and assume that $m_n = \lambda^n$, $n \in \mathbb{N}$. The unbiased case corresponds to $\lambda = 1$, so without loss of generality we assume that $\lambda \neq 1$. The distribution of μ when $\lambda > 1$ is exactly the same as the distribution of μ when the parameter is $1/\lambda$, just with the order of the N associated random variables $\mu_1, \dots, \mu_N$ reversed. Both distributions are multinomial with the same parameters. Therefore, without loss of generality, we do the calculations only for $\lambda \in (0, 1)$. Then, $M_N = \sum_{n=1}^N \lambda^n = \lambda(1 - \lambda^N)/(1 - \lambda) \rightarrow \lambda/(1 - \lambda) =: M$ as $N \rightarrow \infty$ and

$$\begin{aligned} \hat{P}_{i,1} &= \frac{1}{M_N^i} \sum_{n=1}^N \lambda^{in} = \left(\frac{1 - \lambda}{\lambda(1 - \lambda^N)} \right)^i \frac{\lambda^i(1 - \lambda^{iN})}{1 - \lambda^i} \\ (31) \quad &= \frac{(1 - \lambda)^i}{1 - \lambda^i} \frac{1 - \lambda^{iN}}{(1 - \lambda^N)^i} \sim \frac{(1 - \lambda)^i}{1 - \lambda^i}. \end{aligned}$$

In particular,

$$c_N = \hat{P}_{2,1} = \frac{(1 - \lambda)^2}{1 - \lambda^2} \frac{1 - \lambda^{2N}}{(1 - \lambda^N)^2} \sim \frac{1 - \lambda}{1 + \lambda}.$$

For arbitrary $i, j \in \mathbb{N}$ it follows similarly as in the previous Example 5.1 (vii) that the limiting formula (30) for $\hat{P}_{i,j}$ holds, but in the formula (29) for $\phi_j(l_1, \dots, l_j)$, the parameter p_n has to be replaced by $p_n := \lambda^n/M = (1 - \lambda)\lambda^{n-1}$, $n \in \mathbb{N}$. For instance, the diagonal entries $\hat{P}_{i,i}$ of the matrix $\hat{P}$ satisfy

$$\begin{aligned} \lim_{N \rightarrow \infty} \hat{P}_{i,i} &= \phi_i(1, \dots, 1) = \sum_{\substack{n_1, \dots, n_i \in \mathbb{N} \\ \text{all distinct}}} p_{n_1} \cdots p_{n_i} \\ &= \left(\frac{1 - \lambda}{\lambda} \right)^i \sum_{\substack{n_1, \dots, n_i \in \mathbb{N} \\ \text{all distinct}}} \lambda^{n_1 + \dots + n_i} = \left(\frac{1 - \lambda}{\lambda} \right)^i \sum_{n \geq i(i+1)/2} \lambda^n A(n, i), \end{aligned}$$

where $A(n, i)$ is the number of vectors $(n_1, \dots, n_i) \in \mathbb{N}^i$ with pairwise distinct components satisfying $n_1 + \dots + n_i = n$. This example essentially coincides with Example 5.1 (vii). Again (see [24, Theorem 2.1]) the model is the domain of attraction (in the sense of Definition 2.1 b)) of a discrete-time Ξ -coalescent, where the measure Ξ assigns its total mass $\Xi(\Delta) = (p, p) = (1 - \lambda)/(1 + \lambda)$ to the single point $p := (p_1, p_2, \dots) \in \Delta$. The same result holds for $\lambda > 1$, but (see the comment at the beginning of this example) in the definition of p_n above, the parameter λ has to be replaced by $1/\lambda$.

6. SKEWED DIRICHLET MODEL

Consider the model where ξ_n has a negative binomial distribution with pgf $f_n(x) = (p/(1-qx))^{a_n}$, where $a_n > 0$, $p \in (0, 1)$, $q := 1 - p$. Note that $m_n := \mathbb{E}(\xi_n) = (q/p)a_n$, $n \in \mathbb{N}$, and that this model belongs to the skewed compound Poisson class (13) with $\theta_n := a_n$, $\phi(z) := -\log(1-z) = \sum_{m=1}^{\infty} z^m/m$, and $z := q$. In this case μ has the Dirichlet multinomial distribution

$$\mathbb{P}(\mu = k) = \frac{\binom{a_1+k_1-1}{k_1} \cdots \binom{a_N+k_N-1}{k_N}}{\binom{A_N+N-1}{N}} = \frac{N!}{k_1! \cdots k_N!} \frac{[a_1]_{k_1} \cdots [a_N]_{k_N}}{[A_N]_N},$$

$k = (k_1, \dots, k_N) \in \Delta(N)$, where $A_N := a_1 + \cdots + a_N$. In other words, the distribution of μ is multinomial with parameters N and $p_1, \dots, p_N$, where the parameters $p_1, \dots, p_N$ are random and Dirichlet distributed with parameters $a_1, \dots, a_N$. The symmetric Dirichlet model is for example studied in [29, p. 847]. From (11), it follows that the forward process X has transition probabilities

$$P_{i,j} = \frac{1}{\binom{N}{i}} \sum_{1 \leq n_1 < \cdots < n_j \leq N} \frac{\binom{a_{n_1} + \cdots + a_{n_j}}{j} \binom{A_N - (a_{n_1} + \cdots + a_{n_j})}{N-j}}{\binom{A_N+N-1}{N}}, \quad i, j \in S.$$

Note that μ has joint descending factorial moments

$$\mathbb{E}((\mu_1)_{l_1} \cdots (\mu_N)_{l_N}) = \frac{(N)_l}{[A_N]_l} \prod_{n=1}^N [a_n]_{l_n}, \quad l_1, \dots, l_N \in \mathbb{N}_0,$$

where $l := l_1 + \cdots + l_N$. In particular, $\mathbb{E}(\mu_n) = Na_n/A_N$, $n \in \{1, \dots, N\}$. We conclude from (12) that the backward process $\hat{X}$ has transition probabilities

$$\hat{P}_{i,j} = \frac{i!}{[A_N]_i} \sum_{1 \leq n_1 < \cdots < n_j \leq N} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \cdots + l_j = i}} \prod_{k=1}^j \frac{[a_{n_k}]_{l_k}}{l_k!}, \quad i, j \in S,$$

with the convention that $\hat{P}_{i,0} = \delta_{i0}$. In particular, $\hat{P}_{i,1} = ([A_N]_i)^{-1} \sum_{n=1}^N [a_n]_i$, $i \in \{1, \dots, N\}$, and hence, $c_N = \hat{P}_{2,1} = ([A_N]_2)^{-1} \sum_{n=1}^N [a_n]_2 > 0$. From $a_n/A_N \leq (a_n + 1)/(A_N + 1)$ it follows that $c_N \geq \sum_{n=1}^N (a_n/A_N)^2 \geq 1/N$, or, equivalently, $N_e = 1/c_N \leq N$. Moreover, $d_N = \hat{P}_{3,1} = ([A_N]_3)^{-1} \sum_{n=1}^N [a_n]_3 > 0$.

It is now verified that $c_N \rightarrow 0$ if and only if $d_N/c_N \rightarrow 0$. The basic idea of the proof is the same as for the skewed Kimura model, however, the technical details are a bit more involved. We have

$$\max_{1 \leq n \leq N} \left(\frac{a_n}{A_N} \right)^2 \leq \sum_{n=1}^N \left(\frac{a_n}{A_N} \right)^2 \leq c_N.$$

Thus, $\max_{1 \leq n \leq N} (a_n/A_N) \leq c_N^{1/2}$. Moreover,

$$\begin{aligned} d_N &= \sum_{n=1}^N \frac{a_n(a_n+1)(a_n+2)}{A_N(A_N+1)(A_N+2)} \leq \sum_{n=1}^N \frac{a_n(a_n+1)}{A_N(A_N+1)} \left(\frac{a_n}{A_N} + \frac{2}{A_N} \right) \\ &\leq \left(\max_{1 \leq n \leq N} \frac{a_n}{A_N} \right) \sum_{n=1}^N \frac{a_n(a_n+1)}{A_N(A_N+1)} + \frac{2}{A_N} \sum_{n=1}^N \frac{a_n(a_n+1)}{A_N(A_N+1)} \\ &= \left(\max_{1 \leq n \leq N} \frac{a_n}{A_N} \right) c_N + \frac{2}{A_N} c_N \leq c_N^{3/2} + \frac{2}{A_N} c_N. \end{aligned}$$

Thus, $d_N/c_N \leq c_N^{1/2} + 2/A_N$. Suppose now that $c_N \rightarrow 0$. From the formula for c_N it is readily seen that $c_N \geq a_1^2/[A_N]_2$. Thus, $[A_N]_2 \geq a_1^2/c_N \rightarrow \infty$. It follows that $A_N \rightarrow \infty$ and, hence, $d_N/c_N \leq c_N^{1/2} + 2/A_N \rightarrow 0$. Conversely, if $d_N/c_N \rightarrow 0$, then $c_N \rightarrow 0$ (see Section 2). Thus, $c_N \rightarrow 0$ if and only if $d_N/c_N \rightarrow 0$.

Example 6.1. (power law growth) Suppose that $a_n := n^{-\alpha}$ for some constant $\alpha \in \mathbb{R}$. Five ranges for the parameter α are distinguished. For $\alpha \in (-\infty, 0)$, $A_N \sim N^{1-\alpha}/(1-\alpha) \rightarrow \infty$,

$$c_N \sim \frac{1}{A_N^2} \sum_{n=1}^N a_n^2 \sim \frac{(1-\alpha)^2}{1-2\alpha} \frac{1}{N} \quad \text{and} \quad d_N \sim \frac{1}{A_N^3} \sum_{n=1}^N a_n^3 \sim \frac{(1-\alpha)^3}{1-3\alpha} \frac{1}{N^2}.$$

For $\alpha = 0$ we have $c_N = 2/(N+1) \sim 2/N$ and $d_N = 6/((N+1)(N+2)) \sim 6/N^2$. If $\alpha \in (0, 1)$, then $c_N \sim 1/A_N \sim (1-\alpha)/N^{1-\alpha}$ and $d_N \sim 2/A_N^2 \sim 2(1-\alpha)^2/N^{2-2\alpha}$. If $\alpha = 1$, then $c_N \sim 1/\log N \rightarrow 0$ and $d_N \sim 2/\log^2 N$. In all these four cases considered so far we have $c_N \rightarrow 0$ and $d_N/c_N \rightarrow 0$, so the model is in the domain of attraction of the Kingman coalescent. Suppose now that $\alpha \in (1, \infty)$. Then, $A_N = \sum_{n=1}^N n^{-\alpha} \rightarrow A := \zeta(\alpha) > 1$, $\sum_{n=1}^N a_n(a_n+1) = \sum_{n=1}^N (n^{-2\alpha} + n^{-\alpha}) \rightarrow \zeta(2\alpha) + \zeta(\alpha)$, and, hence, $c_N \rightarrow (\zeta(2\alpha) + \zeta(\alpha))/[\zeta(\alpha)]_2 > 0$. Similarly, it follows that $d_N \rightarrow (\zeta(3\alpha) + 3\zeta(2\alpha) + 2\zeta(\alpha))/[\zeta(\alpha)]_3 > 0$.

For $l_1, \dots, l_j \in \mathbb{N} \setminus \{1\}$ and $N \geq i := l_1 + \dots + l_j$,

$$\begin{aligned} \Phi_j(l_1, \dots, l_j) &= \frac{1}{[A_N]_i} \sum_{\substack{n_1, \dots, n_j=1 \\ \text{all distinct}}}^N [a_{n_1}]_{l_1} \cdots [a_{n_j}]_{l_j} \\ &\rightarrow \frac{1}{[A]_i} \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} [a_{n_1}]_{l_1} \cdots [a_{n_j}]_{l_j} = \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} \mathbb{E}(D_{n_1}^{l_1} \cdots D_{n_j}^{l_j}), \end{aligned}$$

where $(D_{n_1}, \dots, D_{n_j}, 1 - \sum_{k=1}^j D_{n_k})$ has a Dirichlet distribution with parameters $a_{n_1}, \dots, a_{n_j}, A - \sum_{k=1}^j a_{n_k}$. Thus, the limit $\phi_j(l_1, \dots, l_j) := \lim_{N \rightarrow \infty} \Phi_j(l_1, \dots, l_j)$ exists and is of the form $\phi_j(l_1, \dots, l_j) = \int_{\Delta_j} x_1^{l_1} \cdots x_j^{l_j} M_j(dx_1, \dots, dx_j)$, where $\Delta_j := \{(x_1, \dots, x_j) \in [0, 1]^j : x_1 + \dots + x_j \leq 1\}$, and the symmetric measure

$$(32) \quad M_j := \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} P_{n_1, \dots, n_j}$$

on Δ_j is an infinite mixture of the distributions $P_{n_1, \dots, n_j}$ of $(D_{n_1}, \dots, D_{n_j})$. The measures $M_1, M_2, \dots$ completely characterize the limiting discrete-time coalescent process and the model is in the domain of attraction of this coalescent. We now

determine the characterizing measure Ξ of this coalescent. For $j \in \mathbb{N}$ let Q_j denote the probability measure on Δ_j with density

$$(33) \quad f_j(x) := \frac{\Gamma(A)}{\Gamma(a_1) \cdots \Gamma(a_j) \Gamma(A - A_j)} x_1^{a_1-1} \cdots x_j^{a_j-1} \left(1 - \sum_{k=1}^j x_k\right)^{A-A_j-1},$$

$x = (x_1, \dots, x_j) \in \Delta_j$, with respect to the Lebesgue measure on Δ_j . Liouville's integration formula shows that $\int_0^{1-(x_1+\dots+x_j)} f_{j+1}(x_1, \dots, x_{j+1}) dx_{j+1} = f_j(x_1, \dots, x_j)$ for $j \in \mathbb{N}$ and $(x_1, \dots, x_j) \in \Delta_j$, so the sequence $(Q_j)_{j \in \mathbb{N}}$ is consistent. Kolmogoroff's extension theorem ensures that there exists a probability measure Q on $\mathbb{R}^{\mathbb{N}}$ such that, for each $j \in \mathbb{N}$, the projection of Q on $\mathbb{R}^j$ coincides with Q_j . The probability measure Q is called (see, for example, [11, Section 4]) the Dirichlet-Kingman distribution with parameter $(a_n)_{n \in \mathbb{N}}$. Let $D_1, D_2, \dots$ be random variables with joint distribution Q and let $D_{(1)} \geq D_{(2)} \geq \dots$ denote the $D_1, D_2, \dots$ in decreasing order. Note that D_n has a beta distribution with parameters a_n and $A - a_n$, $n \in \mathbb{N}$. In particular, $\mathbb{E}(D_n) = a_n/A$ for $n \in \mathbb{N}$. Now let ν denote the joint distribution of $(D_{(1)}, D_{(2)}, \dots)$ and define $\Xi(dx) := (x, x)\nu(dx)$. Then, for $j, l_1, \dots, l_j \in \mathbb{N}$,

$$\begin{aligned} \int_{\Delta} \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} x_{n_1}^{l_1} \cdots x_{n_j}^{l_j} \frac{\Xi(dx)}{(x, x)} &= \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} \mathbb{E}(D_{(n_1)}^{l_1} \cdots D_{(n_j)}^{l_j}) \\ &= \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} \mathbb{E}(D_{n_1}^{l_1} \cdots D_{n_j}^{l_j}) = \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} \frac{[a_{n_1}]_{l_1} \cdots [a_{n_j}]_{l_j}}{[A]_{l_1 + \dots + l_j}} = \phi_j(l_1, \dots, l_j), \end{aligned}$$

which shows that Ξ is the characterizing measure of this coalescent. Note that for all M_j -integrable functions g it follows that $\mathbb{E}(\sum_{n_1, \dots, n_j} g(D_{n_1}, \dots, D_{n_j})) = \int_{\Delta_j} g dM_j$, where the sum $\sum_{n_1, \dots, n_j}$ extends over all pairwise distinct $n_1, \dots, n_j \in \mathbb{N}$. Hence, the measure M_j , defined in (32), is the j th correlation measure (see, for example, Handa [10, Eq. (2.1)]) of the point process $\sum_{n=1}^{\infty} \delta_{D_n}$.

Example 6.2. (geometric growth) Suppose that $a_n = \lambda^n$ for some $\lambda \in (0, \infty)$. The unbiased case corresponds to $\lambda = 1$.

(i) If $\lambda < 1$, then $A_N \rightarrow A := \lambda/(1 - \lambda)$, so A_N converges. We are hence essentially in the situation of Example 6.1 with $\alpha > 1$. All results there are valid, but now with $A = \lambda/(1 - \lambda)$ (instead of $A = \zeta(\alpha)$) and with $a_n = \lambda^n$ (instead of $a_n = n^\alpha$). In particular, the correlation measures $M_1, M_2, \dots$ of the limiting coalescent process with simultaneous multiple collisions are again infinite mixtures of the form (32).

(ii) If $\lambda > 1$, then $A_N = \sum_{n=1}^N \lambda^n \sim \lambda^{N+1}/(\lambda - 1)$, so A_N does not converge. We are hence essentially in the situation of Example 5.2 with $\lambda > 1$, and it follows that the measure Ξ of the limiting Ξ -coalescent assigns its total mass $\Xi(\Delta) = (\lambda - 1)/(\lambda + 1)$ to the single point $p = (p_1, p_2, \dots) \in \Delta$ defined via $p_n := (\lambda - 1)/\lambda^n$, $n \in \mathbb{N}$.

Remark. (Dirichlet-Kingman coalescent) Examples 6.1 and 6.2 (i) lead to the following more general example of a Ξ -coalescent. Let $a = (a_n)_{n \in \mathbb{N}}$ be a convergent sequence of positive real numbers. Define $A_j := \sum_{n=1}^j a_n$, $j \in \mathbb{N}$, $A := \lim_{j \rightarrow \infty} A_j = \sum_{n=1}^{\infty} a_n$, and, for $j \in \mathbb{N}$, define the function $\phi_j : \mathbb{N}^j \rightarrow \mathbb{R}$ via

$$(34) \quad \phi_j(l_1, \dots, l_j) := \frac{1}{[A]_{l_1 + \dots + l_j}} \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} [a_{n_1}]_{l_1} \cdots [a_{n_j}]_{l_j}.$$

Note that $\phi_1(1) = 1$ and that the functions $\phi_j, j \in \mathbb{N}$, are consistent in the sense that $\phi_j(l_1, \dots, l_j) = \phi_{j+1}(l_1, \dots, l_j, 1) + \sum_{i=1}^j \phi_j(l_1, \dots, l_{i-1}, l_i + 1, l_{i+1}, \dots, l_j)$ for all $j, l_1, \dots, l_j \in \mathbb{N}$. Thus, there exists a discrete-time Ξ -coalescent $R = (R_k)_{k \in \mathbb{N}_0}$ such that, if R is in a state with $l \in \mathbb{N}$ blocks, any $(l_1, \dots, l_j)$ -collision ($j, l_1, \dots, l_j \in \mathbb{N}$ with $l_1 + \dots + l_j = l$) occurs with probability (34). The measure Ξ is obtained as described in Example 6.1. More precisely, $\nu(dx) := \Xi(dx)/(x, x)$ is the joint distribution of the ordered random variables $D_{(1)} \geq D_{(2)} \geq \dots$, where the random variables $D_1, D_2, \dots$ are such that, for each $j \in \mathbb{N}$, $(D_1, \dots, D_j)$ has density (33). The random partition $(D_1, D_2, \dots)$ of the unit interval is (see, for example, [11, Section 4]) the Dirichlet-Kingman partition with parameter a . One may therefore call R the discrete-time Dirichlet-Kingman coalescent with parameter a . A standard method for example well described in [7, p. 163, Eq. (2.3)] allows to transform the discrete-time coalescent R into a continuous-time coalescent $\Pi = (\Pi_t)_{t \geq 0}$ such that, if this coalescent is in a state with $l \in \mathbb{N}$ blocks, any $(l_1, \dots, l_j)$ -collision ($j, l_1, \dots, l_j \in \mathbb{N}$ with $l_1 + \dots + l_j = l > j$) occurs at rate (34). Note however that Π has total rates $1 - \phi_j(1, \dots, 1), j \in \mathbb{N}$. One may call Π the continuous-time Dirichlet-Kingman coalescent with parameter a . This coalescent has many properties in common with the Poisson-Dirichlet coalescent [22]. For example, both measures ν and Ξ are concentrated on the subset Δ^* of points $x \in \Delta$ satisfying $|x| := \sum_{n=1}^{\infty} x_n = 1$. In particular, $\int |x| \nu(dx) = \nu(\Delta) = 1 < \infty$, showing that these coalescents, if started in the partition of $\mathbb{N}$ into singletons, do not have proper frequencies and, hence, cannot come down from infinity.

7. SKEWED KIMURA MODEL

In this section an example is presented which (in contrast to the models in Sections 5 and 6) does not belong to the compound Poisson class. Suppose that for each $n \in \mathbb{N}$ the random variable ξ_n has a binomial distribution with pgf $f_n(x) = (px + q)^{a_n}$, where $a_n \in \mathbb{N}$, $p \in (0, 1)$ and $q := 1 - p$. Note that $m_n := \mathbb{E}(\xi_n) = pa_n, n \in \mathbb{N}$. Then, $\mu = (\mu_1, \dots, \mu_N)$ has a multi-hypergeometric distribution of the form $\mathbb{P}(\mu = k) = \binom{a_1}{k_1} \dots \binom{a_N}{k_N} / \binom{A_N}{N}$, $k = (k_1, \dots, k_N) \in \Delta(N)$, where $A_N := a_1 + \dots + a_N$. From (11), it follows that the forward process X has transition probabilities

$$P_{i,j} = \frac{1}{\binom{N}{i}} \sum_{1 \leq n_1 < \dots < n_i \leq N} \frac{\binom{a_{n_1} + \dots + a_{n_i}}{j} \binom{A_N - (a_{n_1} + \dots + a_{n_i})}{N-j}}{\binom{A_N}{N}}, \quad i, j \in S.$$

Note that μ has joint factorial moments

$$\mathbb{E} \left(\binom{\mu_1}{l_1} \dots \binom{\mu_N}{l_N} \right) = \frac{\binom{A_N}{l_1 + \dots + l_N}}{\binom{A_N}{l_1 + \dots + l_N}} \binom{a_1}{l_1} \dots \binom{a_N}{l_N}, \quad l_1, \dots, l_N \in \mathbb{N}_0,$$

In particular, $\mathbb{E}(\mu_n) = Na_n/A_N, n \in \{1, \dots, N\}$. From (12), it follows that the backward process $\hat{X}$ has transition probabilities

$$\hat{P}_{i,j} = \frac{1}{\binom{A_N}{i}} \sum_{1 \leq n_1 < \dots < n_j \leq N} \sum_{\substack{l_1, \dots, l_j \in \mathbb{N} \\ l_1 + \dots + l_j = i}} \binom{a_{n_1}}{l_1} \dots \binom{a_{n_j}}{l_j}, \quad i, j \in S,$$

with the convention that $\hat{P}_{i,0} = \delta_{i,0}, i \in S$. Note that $\hat{P}_{i,1} = ((A_N)_i)^{-1} \sum_{n=1}^N (a_n)_i, i \in \{1, \dots, N\}$, and, hence, $c_N = \hat{P}_{2,1} = ((A_N)_2)^{-1} \sum_{n=1}^N (a_n)_2$ and $d_N = \hat{P}_{3,1} =$

$((A_N)_3)^{-1} \sum_{n=1}^N (a_n)_3$. The coalescence probability c_N can be smaller than $1/N$ (choose for example $a_n = 2$ for all $n \in \mathbb{N}$ leading to $c_N = 1/(2N - 1)$), so the effective population size $N_e = 1/c_N$ can be larger than N . We have $c_N = 0$ if and only if $a_n = 1$ for all $n \in \{1, \dots, N\}$. For the rest of this section it is assumed that $a_n > 1$ for at least one index $n \in \mathbb{N}$ such that $c_N > 0$ for sufficiently large N . In the following it is verified that $c_N \rightarrow 0$ if and only if $d_N/c_N \rightarrow 0$. From $(a_n - 2)/(A_N - 2) \leq (a_n - 1)/(A_N - 1) \leq a_n/A_N$ it follows that

$$\max_{1 \leq n \leq N} \left(\frac{a_n - 2}{A_N - 2} \right)^2 \leq \max_{1 \leq n \leq N} \frac{a_n(a_n - 1)}{A_N(A_N - 1)} \leq \sum_{n=1}^N \frac{a_n(a_n - 1)}{A_N(A_N - 1)} = c_N.$$

Thus, $\max_{1 \leq n \leq N} ((a_n - 2)/(A_N - 2)) \leq c_N^{1/2}$, and consequently

$$\begin{aligned} d_N &= \sum_{n=1}^N \frac{a_n(a_n - 1)(a_n - 2)}{A_N(A_N - 1)(A_N - 2)} \\ &\leq \left(\max_{1 \leq n \leq N} \frac{a_n - 2}{A_N - 2} \right) \sum_{n=1}^N \frac{a_n(a_n - 1)}{A_N(A_N - 1)} = \left(\max_{1 \leq n \leq N} \frac{a_n - 2}{A_N - 2} \right) c_N \leq c_N^{3/2}, \end{aligned}$$

or, equivalently, $d_N/c_N \leq c_N^{1/2}$. Therefore, if $c_N \rightarrow 0$, then $d_N/c_N \rightarrow 0$. As already mentioned in Section 2, the converse holds for arbitrary Cannings models, so if $d_N/c_N \rightarrow 0$, then $c_N \rightarrow 0$. In particular, the model is in the domain of attraction of the Kingman coalescent if and only if $c_N \rightarrow 0$ as $N \rightarrow \infty$.

Example 7.1. (power law growth) Fix $K \in \mathbb{N}_0$ and suppose that $a_n = n^K$, $n \in \mathbb{N}$. Then, $A_N = \sum_{n=1}^N n^K \sim N^{K+1}/(K+1)$ and, hence,

$$c_N = \frac{1}{(A_N)_2} \sum_{n=1}^N (n^K)_2 \sim \frac{1}{A_N^2} \sum_{n=1}^N n^{2K} \sim \left(\frac{K+1}{N^{K+1}} \right)^2 \frac{N^{2K+1}}{2K+1} = \frac{(K+1)^2}{2K+1} \frac{1}{N}.$$

Thus $N_e = \rho N$ with $\rho := (2K+1)/(K+1)^2 < 1$. Similarly,

$$d_N \sim \frac{1}{A_N^3} \sum_{n=1}^N n^{3K} \sim \left(\frac{K+1}{N^{K+1}} \right)^3 \frac{N^{3K+1}}{3K+1} = \frac{(K+1)^3}{3K+1} \frac{1}{N^2}.$$

In particular, $c_N \rightarrow 0$, so the model is in the domain of attraction of the Kingman coalescent. This example essentially coincides with Example 5.1 (i) with $\alpha := -K$.

Example 7.2. (geometric growth) Fix $\lambda \in \{2, 3, \dots\}$ and suppose that $a_n = \lambda^n$, $n \in \mathbb{N}$. Then, $A_N = \sum_{n=1}^N \lambda^n = \lambda(\lambda^N - 1)/(\lambda - 1) \sim \lambda^{N+1}/(\lambda - 1)$. Now proceed as in Example 5.2 with $\lambda > 1$. Define $b := 1/\lambda$ and $p_n := (1 - b)b^{n-1} = (\lambda - 1)/\lambda^n$, $n \in \mathbb{N}$. For $l_1, \dots, l_j \in \mathbb{N}$, it follows that

$$\begin{aligned} \Phi_j(l_1, \dots, l_j) &= \frac{1}{(A_N)_j} \sum_{\substack{n_1, \dots, n_j=1 \\ \text{all distinct}}}^N (a_{n_1})_{l_1} \cdots (a_{n_j})_{l_j} \\ &\sim \frac{1}{A_N^j} \sum_{\substack{n_1, \dots, n_j=1 \\ \text{all distinct}}}^N \lambda^{n_1 l_1 + \dots + n_j l_j} \rightarrow \sum_{\substack{n_1, \dots, n_j \in \mathbb{N} \\ \text{all distinct}}} p_{n_1}^{l_1} \cdots p_{n_j}^{l_j}, \end{aligned}$$

where the proof of the last convergence works as in Example 5.2. The measure Ξ of the limiting Ξ -coalescent assigns its total mass $\Xi(\Delta) = (\lambda - 1)/(\lambda + 1)$ to the single point $p = (p_1, p_2, \dots) \in \Delta$.

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