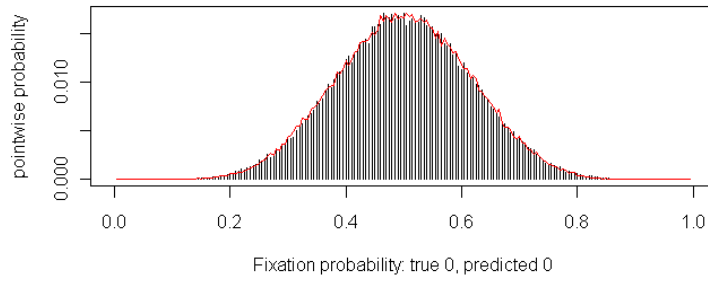
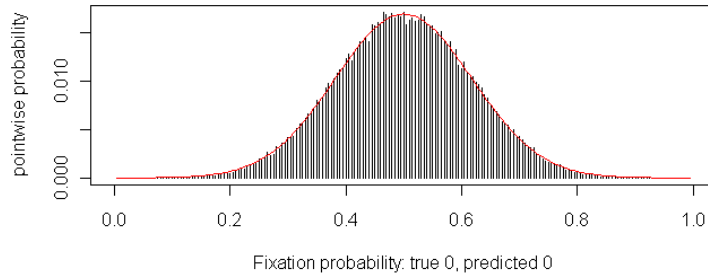


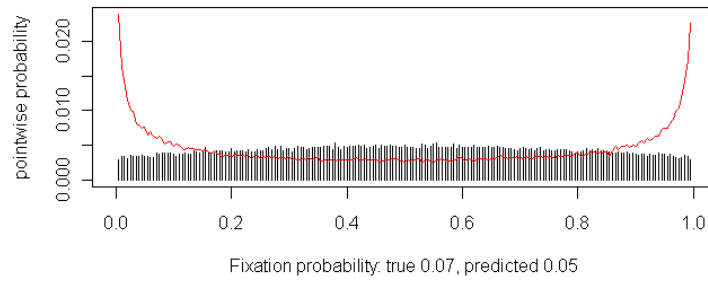
Multinomial-Dirichlet vs. true distribution, scenario 1



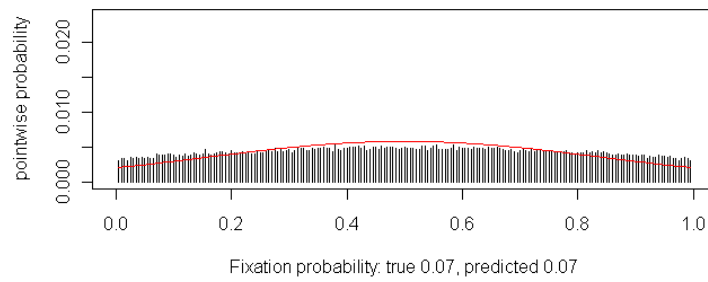
Truncated normal vs. true distribution, scenario 1



Multinomial-Dirichlet vs. true distribution, scenario 2



Truncated normal vs. true distribution, scenario 2



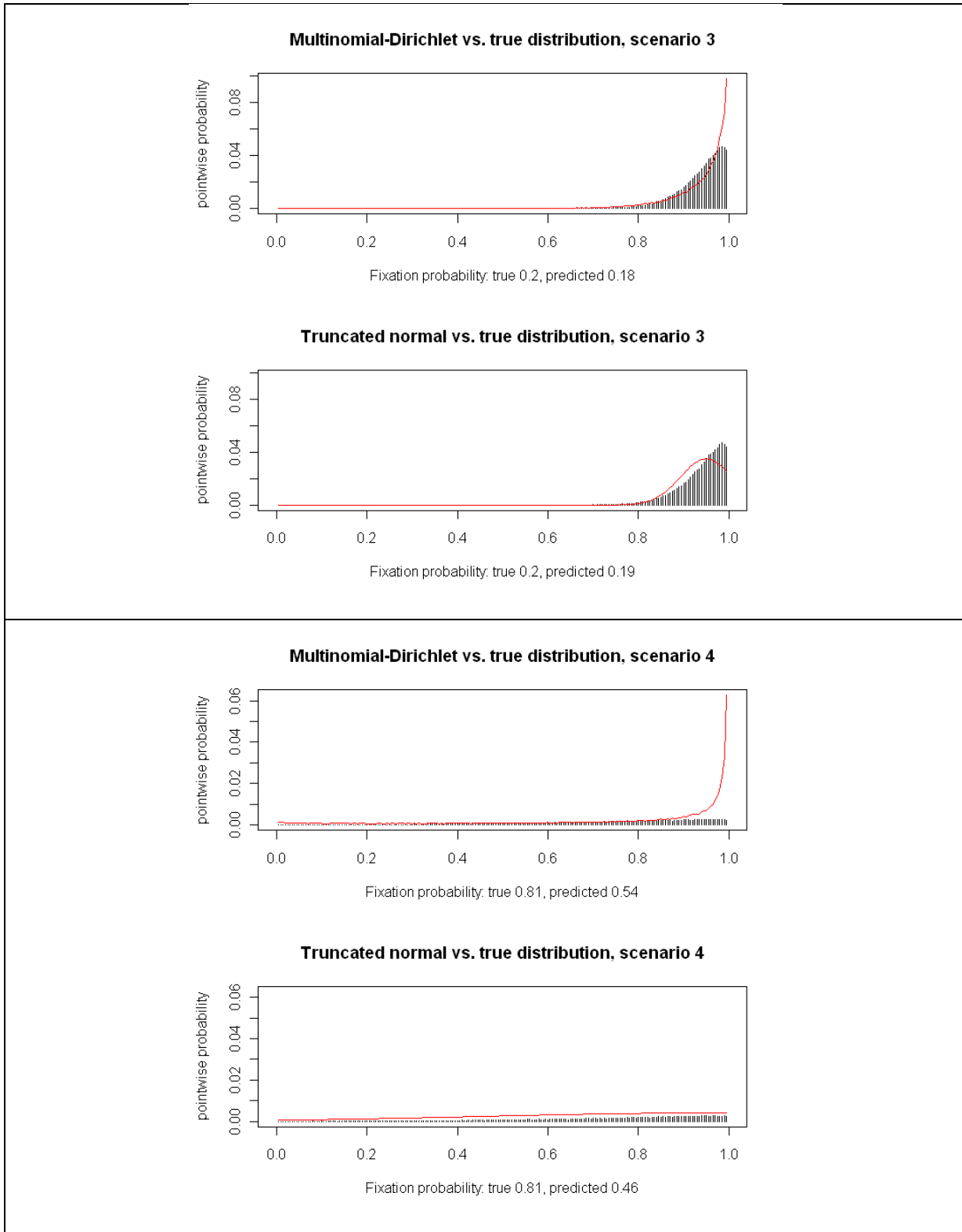


Figure S1 This figure represents an empirical sample from the true distribution of allele frequency (black discrete distribution) in four scenarios, and two approximations to it: Multinomial-Dirichlet and truncated normal. The parameter values are: Scenario 1: $T = 10$ generations, $n_A = 100$ individuals, initial frequency $q_{j1} = 0.5$; Scenario 2: $T = 100$, $n_A = 100$, $q_{j1} = 0.5$; Scenario 3: $T = 10$, $n_A = 100$, $q_{j1} = 0.95$; Scenario 4: $T = 100$, $n_A = 100$, $q_{j1} = 0.95$.

File S1

Calculating coancestry coefficients

Calculating coancestry coefficients from the admixture F-model (AFM). The frequency of allele u in subpopulation A is simply the average of the indicator variables x_{ijk_u} ,

$$p_{Aju} = \frac{1}{2n_A} \sum_{i=1}^{n_A} (x_{ij1u} + x_{ij2u}). \quad (\text{Eq. S1})$$

The frequency p_{Aju} is a random variable, with expectation (over the flow of neutral alleles through a pedigree structure that we consider fixed) q_{ju} . The covariance among subpopulations A and B is

$$\text{Cov}(p_{Aju}, p_{Bju}) = E[(p_{Aju} - p_{ju})(p_{Bju} - p_{ju})] = E(p_{Aju}p_{Bju}) - q_{ju}^2 = (q_{ju} - q_{ju}^2)\theta_{AB}^p, \quad (\text{Eq. S2})$$

where the last equality follows (at the limit of low mutation rate) by noting that identity by state follows either from identity by descent, or by the two distinct alleles in the ancestral population being identical by state,

$$E(x_{i'ju}x_{iju}) = q_{ju}^2 + \theta_{ii'}(q_{ju} - q_{ju}^2) \quad (\text{Eq. S3})$$

and substituting the definitions of p_{Aju} (Eq. S1) and θ_{AB}^p (Eq. 1 in the main text) into Equation S1.

In the AFM, it holds that

$$\text{Cov}(p_{Aju}, p_{Bju}) = \text{Cov}\left(\sum_{k=1}^{n_L} \kappa_{Ak} z_{kju}, \sum_{k=1}^{n_L} \kappa_{Bk} z_{kju}\right). \quad (\text{Eq. S4})$$

Because the lineages are independent, $\text{Cov}(z_{kju}, z_{k'ju})=0$ for all $k \neq k'$. Thus, Equation S4 reduces to

$$\text{Cov}(p_{Aju}, p_{Bju}) = \sum_{k=1}^{n_L} \kappa_{Ak}\kappa_{Bk} \text{Var } z_{kju} = \sum_{k=1}^{n_L} \kappa_{Ak}\kappa_{Bk} \frac{q_{ju} - q_{ju}^2}{a_k + 1}. \quad (\text{Eq. S5})$$

Combining this with Eq. S2 yields

$$(q_{ju} - q_{ju}^2)\theta_{AB}^p = \sum_{k=1}^{n_L} \kappa_{Ak}\kappa_{Bk} \frac{q_{ju} - q_{ju}^2}{a_k + 1}, \quad (\text{Eq. S6})$$

and hence

$$\theta_{AB}^p = \sum_{k=1}^{n_L} \frac{\kappa_{Ak}\kappa_{Bk}}{a_k + 1}. \quad (\text{Eq. 12, main text})$$

Calculating coancestry coefficients from a pedigree. If the complete pedigree is known, it is easy to calculate coancestry coefficients for each pair of individuals using the recursive formula (Lynch and Walsh 1998),

$$\theta_{ii'} = \frac{\theta_{is(i)} + \theta_{id(i')}}{2} = \frac{\theta_{s(i)s(i')} + \theta_{s(i)d(i')} + \theta_{d(i)s(i')} + \theta_{d(i)d(i')}}{4} \text{ for } i \neq i',$$
$$\theta_{ii} = \frac{1 + \theta_{s(i)d(i)}}{2}. \quad (\text{Eq. S7})$$

Above, $s(i)$ and $d(i)$ are the sire and dam of individual i , respectively. We used this formula for calculating the true value of θ^P in our simulated data sets.

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Models for allele frequencies

The vector of allele counts at locus j in an isolated population A for generation $t + 1$ follows the multinomial distribution

$$n_{Aj}(t + 1) | n_{Aj}(t) \sim \text{Mult}(2n_A, p_{Aj}(t)) \quad (\text{Eq. S8})$$

where $p_{Aj}(t) = \frac{n_{Aj}(t)}{2n_A}$ is the allele frequency for the generation t . While n_{Aj} follows a multinomial random walk, p_{Aj} follows a corresponding process on an $n_j - 1$ dimensional simplex. This discrete process is often approximated by a continuous-valued random process, the so-called Wright-Fisher diffusion (see e.g. Nicholson *et al.* 2002). Kimura (1955) first derived the exact solution for the distribution of allele frequencies of a biallelic locus under Wright-Fisher diffusion. This solution is not Gaussian, because the diffusion is non-isotropic. Solutions have also been obtained for multiallelic loci (Tavaré 1984; Xie 2011). However, implementing these solutions in the AFM framework would pose considerable computational challenges because of the need to iterate infinite, high-dimensional sums. In case of biallelic loci, the solution of Wright-Fisher diffusion is often approximated by a truncated normal distribution (e.g. Balding 2003; Coop *et al.* 2010; Nicholson *et al.* 2002). However, this approximation cannot be applied on multiallelic loci as such, because the distribution of p_{Aj} needs to be restricted on the simplex Δ^{n_j-1} . The alternative that we apply here is to use the Dirichlet distribution as a phenomenological, i.e. non-mechanistic, model for allele frequencies.

Application of the Dirichlet distribution as a model of pure drift may be considered questionable for two reasons. Firstly, the Dirichlet distribution is known to arise as an equilibrium distribution from the balance of random drift and mutation or migration (e.g. Nicholson *et al.* 2002; Rannala 1996), but not as a result of pure random drift in an isolated population. Secondly, the Dirichlet distribution is a continuous distribution such that each component is restricted on the open interval $]0,1[$, which gives a zero probability for the fixation of any one allele. However, with a small value of the parameter a_A , the Dirichlet distribution can have much of its probability mass very close to the boundaries. Thus, when supplemented with a sampling model for a finite population,

$$n_{Aj} \sim \text{Mult}(2n_A, p'_{Aj}),$$

$$p'_{Aj} \sim \text{Dirichlet}(a_A q_j),$$

the Dirichlet model is able to predict a high probability of fixation. In the AFM, we use Dirichlet-distributed allele frequencies z_{Aj} to model the evolutionary history of the independent lineages, and the multinomial step naturally follows

from the fact that the sample of genotypes is finite, even if the whole subpopulation is sampled. Below, we investigate this model by a comparison with the truncated normal distribution in a biallelic case where both distributions are easily tractable.

We consider a closed population of N individuals that mates randomly for T generations, and assume that the initial frequency of allele 1 has been q_{j1} . In this Supplement, we focus on four representative cases: symmetric allele frequencies with moderate drift (Scenario 1, Fig. S1), symmetric allele frequencies with a high amount of drift (Scenario 2, Fig. S1), uneven allele frequencies with moderate drift (Scenario 3, Fig. S1) and uneven allele frequencies with a high amount of drift (Scenario 4, Fig. S1). To sample from this model, we first generated a sample of size 10^5 from the last generation by using the true model (repeated application of Eq. S8). Then, we derived a corresponding sample from the Dirichlet approximation by randomizing \mathbf{p}'_{Aj} for 10^5 times and sampling the allele counts for each realization from $\text{Mult}(2n_A, \mathbf{p}'_{Aj})$. Finally, we considered the model of allele frequencies under the truncated normal approximation. As suggested by Nicholson et al. (2002), we specified the allele frequency as

$$p_{Aj1} \sim \text{N}(q_{j1}, cq_{j1}(1 - q_{j1})) := \Phi$$

so that the extinction probability of allele 1 was calculated as $\Phi(0)$ and the fixation probability as $1 - \Phi(1)$. We calculated the pointwise probabilities of the discrete classes as $\Phi'(p_{Aj1})/2n_A$, i.e. by dividing the Gaussian density function by the number of discrete values in $]0,1[$. While theoretical values exist for the drift parameters c and a_A given the demographic model, we optimized the values of these parameters in each scenario by minimizing the square distance (denoted D^2) with the true (empirical) distribution.

The results show that both the Dirichlet and truncated normal are imperfect approximations. In scenario 1, where drift is moderate and fixations do not occur, both approximations are qualitatively good, while the truncated normal distribution has a better goodness of fit (Multinomial-Dirichlet $D^2 = 9.2 \times 10^{-8}$; truncated normal $D^2 = 5.9 \times 10^{-8}$). In scenario 2, the truncated normal approximation has a better goodness of fit ($D^2 = 5.9 \times 10^{-7}$), than the Dirichlet approximation ($D^2 = 1.5 \times 10^{-5}$) which has an inconveniently convex shape in this case. In scenario 3, the truncated normal approximation has a better goodness of fit ($D^2 = 1.1 \times 10^{-5}$), but it is qualitatively different from the data by having a clear mode in the interior of $[0,1]$ which the Dirichlet approximation ($D^2 = 2.5 \times 10^{-5}$) does not have. In scenario 4, the Dirichlet approximation is better ($D^2 = 3.8 \times 10^{-4}$ as opposed to truncated normal $D^2 = 6.1 \times 10^{-4}$). In general, both distributions have problems in coping with the data when the amount of drift is high, which shows in the increase of the square distances. Finally, we note that the expectation of the truncated normal distribution is not strictly q_{j1} which would be expected under pure random drift. On the other hand, this is likely to be unimportant when the amount of drift is low.

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File S3

The MCMC sampling scheme

We used a random-walk Metropolis-Hastings algorithm to sample the joint posterior density of \mathbf{q} , \mathbf{a} and $\boldsymbol{\kappa}$. Below, we describe how each parameter was sampled while keeping the other parameters fixed.

- Sampling the drift parameters \mathbf{a} . We used $N(a_k, \delta_{a_k}^2)$ distributions separately for each k to draw proposals for $\log a_k$. The variance parameters $\delta_{a_k}^2$ were adjusted during the burn-in as in Ovaskainen *et al.* (2008) to give an accept ratio of 0.44.
- Sampling lineage loadings $\boldsymbol{\kappa}$. We used $\text{TDD}(\delta_{\boldsymbol{\kappa}_A}, \boldsymbol{\kappa}_A)$, i.e. truncated Dirichlet, distributions (Fang *et al.* 2000) separately for each A and j to draw proposals for $\boldsymbol{\kappa}_A$. The $\delta_{\boldsymbol{\kappa}_A}$'s are proposal parameters that were adjusted during the burn-in as in Ovaskainen *et al.* (2008) to give an accept ratio of 0.44.
- Sampling ancestral allele frequencies \mathbf{q} and lineage-specific allele frequencies \mathbf{z} . We used $\text{TDD}(\delta_{q_j}, q_j)$ and $\text{TDD}(\delta_{z_{kj}}, z_{kj})$ distributions separately for each j and k to draw proposals for the allele frequencies. The δ_{q_j} 's are proposal parameters that are adjusted during the burn-in as in Ovaskainen *et al.* (2008) to give an accept ratio of 0.44.

We thus used the truncated Dirichlet distribution of Fang *et al.* (2000) to perform the Metropolis-Hastings random walk for the Dirichlet-distributed variables $\boldsymbol{\kappa}$, \mathbf{q} and \mathbf{z} with a pre-set truncation threshold $\tau = 10^{-7}$. This greatly improves the mixing properties of the Markov chain, because it helps to avoid numerical problems on the boundary of the parameter space (i.e. on the edges of the simplices Δ^{n_j-1} and Δ^{n_p-1}). According to our observation, the method that Fang *et al.* (2000) present for sampling from TDD may produce biased samples for high truncation thresholds such as $\tau = 10^{-1}$. However, to our experience, this does not compromise the statistical power of our algorithm with $\tau = 10^{-7}$.

We have implemented the algorithm described above in the R-package RAFM (Karhunen 2012).

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