Title: The effects on neutral variability of recurrent selective sweeps and background selection

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Key words: Selective sweeps, background selection, gene conversion, crossing over, neutral variability, favorable mutations, *Drosophila melanogaster*

Running title: Selective sweeps and background selection

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ABSTRACT

Levels of variability and rates of adaptive evolution can be affected by hitchhiking, the effect of selection on variants at linked sites. Hitchhiking can be caused either by selective sweeps or by background selection, involving the spread of new favorable alleles or the elimination of deleterious mutations, respectively. Recent analyses of population genomic data have fitted models where both these processes act simultaneously, in order to infer the parameters of selection. Here, we investigate the consequences of relaxing a key assumption of some of these studies – that neutral variability at a site affected by recurrent selective sweeps fully recovers between successive sweeps. We derive a simple expression for the expected level of neutral variability in the presence of recurrent selective sweeps and background selection. We also derive approximate integral expressions for the effects of recurrent selective sweeps on a given gene. The accuracy of the theoretical predictions was tested against multilocus simulations, using the software SLiM with selection, recombination and mutation parameters that are realistic for *Drosophila melanogaster*. We find good agreement between the simulation results and predictions from the integral approximations, except when rates of crossing over are close to zero. We show that the observed relations between the rate of crossing over and the level of synonymous site diversity and rate of adaptive evolution largely reflect background selection, whereas selective sweeps are needed to produce substantial distortions of the site frequency spectrum.

The effect of selection at a given locus on the properties of neutral variability at linked sites is a classic problem in population genetics, first studied by Sved (1968) and Ohta and Kimura (1970) in the context of associative overdominance – the apparent heterozygote advantage induced at a neutral locus by variants at linked loci, maintained by heterozygote advantage or by mutation to partially recessive deleterious alleles. This early work was followed by the classic paper of Maynard Smith and Haigh (1974) on the hitchhiking effect, whereby the spread of a favorable mutation reduces the level of neutral variability at a linked locus; this process has come to be termed a 'selective sweep' (Berry *et al.* 1991). It was later shown that selection against recurrent deleterious mutations also reduces neutral variability at linked sites by a hitchhiking process, known as background selection (Charlesworth *et al.* 1993). The same basic equation that describes the effect of selection at one locus on the allele frequency at a linked neutral locus has recently been shown to underlie all three of these processes (Zhao and Charleworth 2016); this expression is an example of the Price-Robertson covariance equation for the effect of selection on one trait on the mean of a correlated trait (Robertson 1968; Price 1970; Santiago and Caballero 1995).

A useful heuristic for viewing these processes was provided by the conceptual framework introduced by Robertson (1961), according to which selection causes a reduction in the effective population size, N_e. This affects both neutral variability and the efficacy of selection (Hill and Robertson 1966), by what is now termed the Hill-Robertson effect (Felsenstein 1974). A large theoretical literature has grown up since these pioneering studies, reviewed by Barton (2010), Stephan (2010), Charlesworth (2012a) and Neher (2013).

Much of the motivation for these theoretical studies came from the advent of data on genome-wide patterns of variability, which inspired attempts to infer the nature and parameters of selection from observations such as the correlation between the level of synonymous sequence diversity in a gene and the local recombination rate (Begun and Aquadro 1992), or its nonsynonymous divergence from a related species (Andolfatto 2007). Work of this type has recently been reviewed by Sella *et al.* (2009), Vitti *et al.* (2013), Booker *et al.* (2017) and Hermisson and Pennings (2017). Several recent studies have used the theory of the joint effects of recurrent selective sweeps and background selection, pioneered by Wiehe and Stephan (1993) and Kim and Stephan (2000), to estimate their effects on levels of neutral diversity across the genomes of multiple species

(Corbett-Detig *et al.* 2015), and to infer the rates of occurrence of advantageous mutations and the strength of selection acting on them (Elyashiv *et al.* 2016; Campos *et al.* 2017). These studies all concluded that the level of variability in a species is often much smaller than would be expected in the absence of selection, even in regions with relatively high rates of genetic recombination. This reduction in variability reflects the effects of both selective sweeps (SSWs) and background selection (BGS), although the estimates of the parameters involved differ substantially among the different studies.

Several important assumptions underlie the model of recurrent sweeps used in this recent work. One is that the effect of BGS on the probability of fixation of a linked favorable mutation is well approximated by its effect on neutral variability at a site at the same location in the genome, which is described by a factor B that multiplies the value of N_e for that site. An expression for B can be found from the standard equation for the effect of BGS in the presence of recombination (Hudson and Kaplan 1995; Nordborg *et al.* 1996). Using the formula of Kimura (1962) for an autosomal, semi-dominant mutation with selective advantage s_A in homozygotes, the probability of fixation of a mutation with $N_e s_A >> 1$ in a randomly mating, discrete-generation population is then $BN_e s_A/N$ instead of $N_e s_A/N$, where N is the population size (Peck 1994; Barton 1995; Stephan *et al.* 1999; Kim and Stephan 2000). This approach breaks down, however, when the product of N_e and the selection coefficient against deleterious mutations is of the order of 1 or less, especially when there is little or no recombination (Gordo *et al.* 2002; Kaiser and Charlesworth 2009; Good *et al.* 2014; Zhao and Charleworth 2016).

The second assumption is that neutral diversity at a site that has experienced a selective sweep will have fully recovered its equilibrium level before the site is affected by another sweep; this also implies that selective sweeps occur sufficiently rarely that they do not interfere with each other. Finally, the theory assumes that sweeps are 'hard', such that each favorable mutation originated as a single copy in the population, as opposed to 'soft' sweeps that arise from standing variation or from several independent mutational events in the same gene (Hermisson and Pennings 2005, 2017).

All of these assumptions can be questioned. The main purpose of this paper is to examine the accuracy of the first two assumptions, in the context of parameter values for BGS and SSWs that appear to be fairly realistic on the basis of inferences from a *Drosophila melanogaster* population (Campos *et al.* 2017). This is done by means of computer simulations of multiple loci that are subject to both BGS and SSWs, together with new approximations for the effects of BGS and SSWs based on replacing summations

across selected sites with integration. The results indicate that there are noticeable deviations from the predictions of models that assume full recovery from one sweep before the occurrence of a new one, consistent with the fact that sweeps occur in single genes at a rate of more than one sweep per $2N_e$ generations. To deal with this problem, we have developed a simple approach to predicting pairwise neutral nucleotide diversity under recurrent selective sweeps, as well as to correct for some other, less important, assumptions. We consider only hard sweeps, because these are amenable to simple analytic modeling and simulation.

Material and Methods

We used the simulation package SLiM (Messer 2013), version 1.8. The details of the simulation methods are described in the online manual (benhaller.com/slim/SLiM.18_manual.pdf). We modeled sets of *n* genes separated by 2kb of selectively neutral intergenic sequence (Figure 1). Each gene had a 5′ UTR and 3′ UTR with length 190 and 280 basepairs, respectively. In addition, there were 5 exons per gene, of length 300bp, each separated by 100bp of neutral intronic sequence. There was thus a total of 2370bp for each gene, which is representative of a typical *Drosophila melanogaster* gene (Campos *et al.* 2017).

In order to simulate realistic parameters of selection, mutation and recombination for a model autosome, we rescaled the values applicable to a natural population of *D. melanogaster* by multiplying them by the ratio of N_e for the population to the number of breeding individuals used in the simulations, *N,* which was usually set to 2500. This conserves the products of N_e and the basic parameters of selection, recombination and mutation. The scaled parameters control most aspects of evolution in finite populations if time is rescaled by a factor of N/N_e (Ewens 2004), such that one generation in the simulations corresponds to N_e/N generations in the natural population. We chose an N_e/N ratio of 532, equivalent to an N_e of 1.33 million for the natural population. This value was based on the mean autosomal synonymous site diversity value of $\pi = 0.018$ for an African population (Campos *et al.* 2017) and a mutation rate of $\mu = 4.5 \times 10^{-9}$ (discussed below), using the standard equilibrium formula for neutral variability under the infinite sites model (Kimura 1971), $\pi = 4N_e\mu$, and assuming (rather conservatively) that diversity has been reduced by hitchhiking effects to 76% of its value in the absence of selection (Campos *et al.* 2017). The

selection, recombination and mutation parameters described below are those considered to be realistic for natural populations of *D. melanogaster*, on the basis of laboratory and population genomic data; the simulation values were obtained by multiplying these by 532.

To model recombination, we mostly used 5 rates of reciprocal crossing over, which were multiples of the standard autosomal recombination rate in Drosophila, adjusted by a factor of ½ to take into account the absence of recombinational exchange in males (Campos *et al.* 2017). These 'effective rates of crossing over' were 0.5×10^{-8}, 1×10^{-8}, 1.5×10^{-8}, 2×10^{-8} and 2.5×10^{-8} cM/Mb, respectively, where 1×10^{-8} is the standard rate. We also ran a limited number of simulations with no crossing over. The simulations were run with and without non-crossover associated gene conversion events, using a rate of initiation of conversion events of 1×10^{-8} cM/Mb and a tract length of 440 bp. These values are consistent with the results of Miller *et al.* (2016), and are similar to estimates from earlier studies of the *rosy* locus (Hilliker *et al.* 1994). We did not vary the rate of initiation of gene conversion when using different rates of crossing over, since this rate appears to be fairly constant across the whole Drosophila genome (Comeron *et al.* 2012; Miller *et al.* 2016).

It should be noted that the procedure of multiplying the female rate of crossing over by a factor of one-half to obtain the parameters used in the autosomal simulations is not entirely realistic, since the simulations assume a fixed map length l for the region in question, and generate numbers of crossovers according to a Poisson distribution with mean l (Haldane 1919). With our assumption, l is one-half the map length in females. The true probability of i crossovers ($i > 1$) is $0.5i^{2l}\exp(-2l)/i!$, as compared to the simulation value of $i^{l}\exp(-l)/i!$, and the true probability of no crossovers is $0.5[1 + \exp(-2l)]$ as compared to $\exp(-l)$. This means that the frequency of crossovers in the region as a whole is considerably over-represented in the simulations. With $l = 1.62$, which corresponds to the case of 70 genes with the standard rate of crossing over, the true probability of no crossovers is 0.520, compared with the simulation value of 0.198. The agreement is somewhat better for the case with one-half the standard value, where the probabilities are 0.599 and 0.445, respectively; as the map length decreases, the two values converge. This problem does not apply to the rate of initiation of gene conversion events, which thus mitigates its effect. In addition, it applies only to recombination between sites separated by relatively large map distances; for local effects of hitchhiking (e.g. within a gene), it is

not a source of serious error. The results described below suggest that most of the hitchhiking effects in our simulations are, in fact, local. While the calculations presented here are not thus strictly realistic as far as representing a *Drosophila* population, this should not invalidate the comparisons between the theoretical predictions and the simulations, which is the main focus of this study.

Mutations were modeled by randomly introducing a new allele at each site, at a rate of 4.5×10^{-9} per site per generation, which is in the midrange of experimental values from DNA sequence analyses of mutation accumulation lines (Schrider *et al.* 2013) and sets of parents and offspring (Keightley *et al.* 2014). We assumed that 30% of exon site mutations were neutral, corresponding to synonymous variants, and the remaining 70% were under selection, corresponding to nonsynonymous (NS) sites. Synonymous pairwise nucleotide site diversity (Nei and Tajima 1983) was used as the measure of neutral variation.

The selection parameters described below are typical of those inferred by Campos *et al.* (2017) for a Rwandan population of *D. melanogaster*, excluding genes in non-crossover genomic regions. The majority of exon site mutations subject to selection were assumed to be deleterious, and to follow a gamma distribution of fitness effects (DFE). The mean scaled selection coefficient for the DFE was $\gamma_{NS} = 2000$ ($\gamma_{NS} = 2N_e s$, where s is the selection coefficient against a homozygous deleterious mutation). The shape parameter of the gamma distribution was set to 0.30. A small fraction of the mutations in NS sites ($p_a = 2.21 \times 10^{-4}$) were assumed to be under positive selection, with a fixed scaled selection coefficient $\gamma_a = 250$. Since SLiM does not separate NS from synonymous sites, we multiplied this value of p_a by a factor of 0.7, so that p_a for a random site in an exon was 1.55×10^{-4}.

All sites in UTRs were assumed to be subject to selection; 5′and 3′UTRs were assigned the same parameters, with most mutations being deleterious, with selective effective following a gamma DFE with shape parameter 0.3 and mean scaled selection coefficient $\gamma_{UT} = 110$, as suggested by the population genomic estimates (Campos *et al.* 2017). A small proportion ($p_u = 9.04 \times 10^{-4}$) of the new mutations were selectively advantageous. Favorable and deleterious mutations at NS and UTR sites were assumed to be semidominant, so that the fitness of heterozygotes at a segregating site was exactly intermediate between the fitnesses of the two homozygotes. For deleterious mutations, selection coefficients (s) greater than 1 (corresponding to homozygous lethality) are set to 1 by SLiM. A large fraction of

deleterious mutations that are lethal when heterozygous ($s \geq 2$) would be undesirable from the point of view of biological reality, since recombination would be ineffective in their heterozygous carriers. Lethality of heterozygous mutations with $N = 2500$ corresponds to a scaled selection coefficient of 10000; for NS mutations with $\gamma_{NS} = 2000$, numerical integration of the gamma distribution with scale parameter 0.3 shows that only 4% of mutations are heterozygous lethal. For UTR mutations, with $\gamma_{UT} = 110$, the fraction of heterozygous lethals is negligible.

In addition to the simulations of autosomes, we ran simulations that were intended to represent X chromosomal mutations with equal fitness effects in the two sexes, but with stronger selection than autosomal mutations, as is expected on both theoretical and empirical grounds (Charlesworth *et al.* 2018). X-linked loci spend two-thirds of their time in females where they can recombine (Campos *et al.* 2013), so that the effective rates of crossing over and initiation of gene conversion events for X-linked loci should be 4/3 times the autosomal values for X-linked genes with similar parameter values in females to the autosomal ones. The version of SLiM that we used did not permit explicit modeling of an X chromosome. We therefore used an autosomal model with a population size of 2500, but assumed that the true N_e was three-quarters of that for the autosomes, as seems to be the case for most X-linked loci with similar effective rates of crossing over to autosomal loci (Campos *et al.* 2013). Because N was kept constant, the autosomal values of the rates of crossing over and initiation of gene conversion events were used in the simulations. In order to ensure that X-linked neutral variability in the absence of selection was three-quarters of the autosomal value, the mutation rate was multiplied by 3/4. Finally, with semi-dominance and equal fitness effects of mutations in males and females, the selection coefficient for an X-linked mutation is 4/3 times that for an autosomal mutation with the same selection coefficient, implying that the scaled selection coefficients are the same ($2N_e s$). To mimic stronger selection for positively selected mutations on the X chromosome, we therefore simply multiplied the scaled selection coefficients by a given factor, either 1.5 or 2. No adjustment was made to the scaled selection coefficient for deleterious mutations. A summary of the parameters used here is given in Table 1.

According to the number of genes simulated, we ran four sets of simulations with genomic regions of 20 (87.4 kb), 70 (305.9 kb), 140 (610 kb) and 210 (920 kb)

genes, each following the gene model shown above. Most of our simulations used multiples of 70 genes because this represents a genomic region with a similar number of genes to the 4th chromosome of *D. melanogaster*, which the simulations with zero crossing over are intended to model. Each simulation was run for 35000 (14*N*) generations, which is amply sufficient to allow the frequency distributions of neutral and deleterious mutations to reach equilibrium (see the online Supplementary Information, File S1, Figure S1). For the final estimates of diversity statistics (nucleotide site diversity, Tajima's *D* and the proportion of singletons) at synonymous, NS, intron and UTR sites) we used data from the final generation of each simulation. For calculating the numbers of fixations of favorable mutations at NS and UTR sites, we recorded the fixations that occurred during the last 20000 (8*N*) generations.

Four different scenarios were simulated. First, purely neutral mutations were simulated in order to calculate the diversity statistics for the neutral reference. Three types of scenario with hitchhiking were simulated (i) SSWs only (ii) BGS only (ii) both SSWs and BGS. Each of these was run with varying numbers of genes, and 20 replicate runs for each model were analyzed. Sample sizes of 20 haploid genomes (a similar size to that used by Campos *et al.* 2017) were used for calculating the population genetic statistics. Mean values of each statistic over genes and replicate runs for a given model were recorded, with upper and lower 2.5 percentiles obtained by bootstrapping the mean values per gene of the chosen statistic across replicates (for brevity, we will refer to these as 95% confidence intervals). The statistics generated by the simulations are presented in the online Supplementary Information, Files S2 and S3.

No new data or reagents were generated by this research. The code for the computer programs used in the models described below is available in the Supplementary Information, File S4.

Theoretical Results

Background selection

The predicted effect of BGS in a multi-site context can be described by the quantity $B = \exp(-E)$, where B is the ratio of expected neutral diversity at a focal neutral site under BGS to its value in the absence of BGS (which is equivalent to the

corresponding ratio of mean coalescence times), and E is the sum of the effects of each selected site (Hudson and Kaplan 1995; Nordborg *et al.* 1996; Santiago and Caballero 1998). We assume a region of chromosome that contains many genes, with selected sites that are continuously distributed with constant density, as in Model 3 of Charlesworth (2012a). We distinguish, however, between nonsynonymous (NS) sites and untranslated regions regions (UTRs). This is, of course, a somewhat crude approximation, given that our genic model includes neutrally evolving intronic and intergenic sequences.

We include both reciprocal exchange via crossing over and non-crossover associated gene conversion in the model. We assume that the main contribution from gene conversion to the effect of recombination on BGS comes from sites that are sufficiently distant that gene conversion causes recombination between them at a fixed rate $g = r_g d_g$ (r_g is the rate of initiation of gene conversion events and d_g is the mean tract length); this is the limiting value of the general expression for the rate of recombination due to gene conversion for sites separated by z basepairs, $g[1 - \exp(-r_g z/d_g)]$ (Frisse *et al.* 2001).

For simplicity, we assume autosomal inheritance, but parallel results hold for X-linked loci, with the appropriate changes in selection, mutation and recombination parameters. Because SLiM assumes no crossover interference, the relation between the frequency of crossing over and map distance in the simulations follows the Haldane mapping function (Haldane 1919), such that the frequency of crossing over between a pair of sites separated by z basepairs is given by:

$$c(z) = \tfrac{1}{2}[1 - \exp(-2r_c z)] \qquad (1)$$

where r_c is the rate of crossing over per basepair.

The net frequency of recombination between the sites is $r(z) = g + c(z)$. The predicted value of E for a given selection coefficient, $t = hs$, against heterozygous carriers of a deleterious mutation, E_t, is given by Equations S1 – S5 in section S1 of the Supplementary Information, File S1. To obtain the final value of E, this equation is numerically integrated over the probability distribution of t values for NS and UTR sites separately, with total deleterious mutation rates U_N and U_U for NS and UTR sites, respectively, giving values E_N and E_U for the corresponding BGS effects.

To mimic the simulation results, we assume a gamma distribution with a shape parameter of 0.3. As in previous studies, we ignore all deleterious mutations with a scaled selection coefficient $\gamma = 2N_e s$ below a critical value γ_c, in order to deal with the problem that very weakly selected mutations are subject to drift and contribute little to BGS effects (Nordborg *et al.* 1996). Following Nordborg *et al.* (1996) and Campos *et al.* (2017), we set $\gamma_c = 5$, and the gamma distributions for both NS and UTR mutations were truncated accordingly. Numerical results for the integral of the kernel of the gamma distribution from γ_c to infinity allow the proportion of mutations that exceed γ_c to be calculated; these are denoted by P_N and P_U for NS and UTR sites, respectively. With the parameters used in the simulations of autosomes, this gives $P_N = 0.871$ and $P_U = 0.694$. The final value for E is given by $P_N E_N + P_U E_U$, from which B can be obtained as $\exp(-E)$.

Selective sweeps

Various methods have been used to predict the approximate effect of a single selective sweep on diversity statistics at a partially linked neutral site in a randomly mating population, as well as for the associated distortion of the neutral site frequency spectrum at segregating sites (Maynard Smith and Haigh 1974; Kaplan *et al.* 1989; Stephan *et al.* 1992; Barton 1998,2000; Gillespie 2000; Durrett and Schweinsberg 2004; Kim 2006; Pfaffelhuber *et al.* 2006; Coop and Ralph 2012; Bossert and Pfaffelhuber 2013). Here we present a simple heuristic derivation of the effect of a sweep on the pairwise nucleotide site diversity, π, based on a combination of coalescent process and diffusion equation approaches. Following earlier approaches, this is done by examining the probability that a neutral lineage that is associated with a favorable allele at the end of a sweep was also associated with it at the start of the sweep, rather with the wild-type allele at the selected locus (Figure 2).

We consider separately the deterministic and stochastic phases of the spread of a favorable mutation, which were identified early in the history of the study of sweeps (Maynard Smith and Haigh 1974; Kaplan *et al.* 1989; Stephan *et al.* 1992; Barton 1998). The initial spread of a favorable allele A_2 from a frequency of $1/(2N)$ is subject to large stochastic effects. With semi-dominance, the probability that A_2 survives this effectively neutral period is approximately $Q = N_e s/N$ in a large population (Kimura 1962), assuming that the scaled selection coefficient, $\gamma = 2N_e s$, is much greater than

one (s is the selective advantage to homozygotes for the favorable mutation). As pointed out by Maynard Smith (1976), the overall expected frequency of A_2 during this quasi-neutral phase (including losses) is approximately $1/(2N)$, after which it starts to behave deterministically. If we condition on the survival of A_2, its expected frequency at the end of the quasi-neutral phase is approximately $1/(2NQ) = \gamma^{-1}$.

In the presence of BGS, we follow Kim and Stephan (2000) and assume that N_e in the formula for fixation probability is multiplied by a constant, B (see above). As will be shown below, this constant may be somewhat different for the effect of BGS on purely neutral processes, such as the recovery of neutral variability from a sweep, and for the effect of BGS on the fixation of favorable mutations. B for the latter is expected to be larger than for the former, since selected variants are more resistant to the effects of hitchhiking (Johnson and Barton 2002). We denote these two constants by B_1 and B_2, respectively, and write λ for the ratio B_1/B_2. The critical frequency at which A_2 can be treated as behaving deterministically is then $(B_2\gamma)^{-1}$. When A_2 reaches a frequency close to 1, there is a second stochastic phase in which it drifts to fixation fairly rapidly, which we consider below. We assume that the other effects of BGS are similar to those for neutral variability, with B_1 as the factor that multiplies N_e.

The expectation of the time spent in the deterministic phase can be found as follows. As described by Ewens (2004, p.169), a semi-dominant favorable allele has the property that the expected time spent in a small interval of allele frequency q to $q + \mathrm{d}q$ is the same as the time spent in the interval $1 - q$ to $1 - q - \mathrm{d}q$. This implies that the expected time that A_2 spends between $1/(2N)$ and $(B_2\gamma)^{-1}$ is the same as the expected time it spends between $1 - (B_2\gamma)^{-1}$ and $1 - 1/(2N)$, so that q during the deterministic phase can conveniently be treated as lying between $(B_2\gamma)^{-1}$ and $1 - (B_2\gamma)^{-1}$. Using the solution of the deterministic selection equation for a semi-dominant allele Haldane (1924), $\mathrm{d}q/\mathrm{d}t = \frac{1}{2}\, spq$, the expected time spent in this interval (expressed in units of coalescent time, $2N_e$ generations) is equal to $2(B_1\gamma)^{-1} \ln(B_2^2\gamma^2) = 4(B_1\gamma)^{-1} \ln(B_2\gamma)$, assuming that the neutral coalescent time is modified by a factor of B_1.

The expected times spent in the two stochastic phases can be found as follows. Using Equation 15 of Ohta and Kimura (1973), and using the fact that N_e is multiplied

by B_1 to take BGS into account, the expected first passage time of a neutral allele from initial frequency $1/(2N)$ to a frequency q is:

$$\bar{T}(q) = 2B_1[(1-q)q^{-1}\ln(1-q)+1] \qquad (2)$$

For $q << 1$, this time is approximately equal to $B_1 q$, so that the additional expected time spent in the first stochastic phase is approximately $\lambda\gamma^{-1}$. By the above symmetry argument, the same applies to the time between $1 - (B_2\gamma)^{-1}$ and $1 - 1/(2N)$. The total expected time to fixation of A_2 when $\gamma >> 1$ is thus:

$$\bar{T}_s \approx 4(B_1\gamma)^{-1}[\ln(B_2\gamma) + \tfrac{1}{2}B_1\lambda] \qquad (3)$$

This expression is very close to Equation A17 of Hermisson and Pennings (2005) for the case with $B_1 = B_2 = 1$, which was derived directly from the diffusion equation for the mean sojourn time of a favorable mutation in a finite population.

As far as the effect of a sweep on neutral diversity is concerned, we note that the rate (in units of coalescent time) at which a neutral lineage that is associated with A_2 at time T recombines onto a background of A_1, conditional on encountering an A_1 haplotype, is $p(T)B_1\rho$, where $p(T)$ is the frequency of the wild-type allele at time T and $\rho = 2N_e r$ is the scaled recombination rate. Here, $T = 0$ at the time of fixation of the favorable allele, and $T = T_s$ at the time when it arose in the population. From the symmetry of the selection equation, the mean frequency of A_1 over this period is 0.5, so that that ρ should be discounted by a factor of ½.

For a single sample path in which A_2 reaches the critical frequency $(B_2\gamma)^{-1}$, the duration of the first stochastic phase is equal to the expected value of the first passage time to this frequency, $\lambda\gamma^{-1}$, plus a random term δT_s with expectation zero and variance $\lambda^2\gamma^{-2}/3$ (File S1, section S2). During this period, a single lineage recombines with A_1 haplotypes at a rate close to $B_1\rho$, since A_1 dominates the population, thus contributing $B_1\rho\,\delta T_s$ to the mean number of recombination events. The final stochastic phase has effectively zero probability of contributing to recombination, due to the prevalence of the favored allele, and can be ignored for this purpose.

The probability P_{cs} that the two sampled haplotypes coalesce as a result of the sweep is equivalent to the probability that neither member of a pair of haplotypes sampled at time $T = 0$ recombined onto an A_1 background, provided that the sweep durations are so short that no coalescence can occur among non-recombined haplotypes during the sweep (Wiehe and Stephan 1993). This probability is given by the first term of a Poisson distribution, whose mean is equal to the expected number of recombination events over the duration of the sweep. We thus have:

$$\begin{aligned} P_{cs} &\approx \mathrm{E}\{\exp[-B_1\rho(\bar{T}_s + 2\delta T_s)]\} \approx \exp(-B_1\rho\bar{T}_s)[1 + \tfrac{1}{2}(2B_1\rho)^2 V_{\delta T_s}] \\ &= \exp\{-4(r/s)[\ln(B_2\gamma) + \tfrac{1}{2}B_1\lambda]\}\ [1 + \tfrac{2}{3}(B_1\lambda r/s)^2] \\ &= (B_2\gamma)^{-4r/s}\exp(-2B_1\lambda r/s)[1 + \tfrac{2}{3}(B_1\lambda r/s)^2] \end{aligned} \qquad (4)$$

The first term on the right-hand side of the third line of Equation 4, corresponding to the deterministic phase contribution, was first derived by Barton (2000) for the case of $B_1 = B_2 = 1$, using a more rigorous approach. It has been used in some subsequent studies (Weissman and Barton 2012; Campos *et al.* 2017). The last term is second-order in r/s and thus is of minor importance, since sweeps only have substantial effects on variability when $r/s << 1$.

Extensions to this result are described in sections S3 and S4 File S1, which allow for the accrual of genetic diversity of swept lineages during the duration of a sweep, and for multiple recombination events that bring a recombined lineage back onto an A_2 background (Figure 2).

Sweeps at multiple sites

We now consider the effects of recurrent sweeps at multiple sites. The standard approach has been to assume that sweeps are sufficiently rare that their effects on a given site can be treated as mutually exclusive events (Wiehe and Stephan 1993), and this assumption will be made here. We consider only a single gene, which is reasonable for favorable mutations whose selection coefficients are less than the rate of recombination between sites in different genes, as is usually the case here. Kim (2006) has derived a general expression that allows for recurrent sweeps, and permits calculations of their effect on the site frequency spectrum at a focal site as well as on

its pairwise diversity, but this does not yield a simple formula of the type that we derive here (Equation 12).

The rate of coalescent events experienced at a given neutral site (in units of $2N_e$ generations), due to recurrent selective sweeps at NS and UTR sites, is then given approximately by:

$$S^{-1} \approx \nu_a \sum_i P_{csN_i} + \nu_u \sum_j P_{csU_j} \qquad (5a)$$

where ν_a and ν_u are the rates (in units of coalescent time) at which substitutions of favorable mutations occur at NS and UTR sites respectively; $P_{cs\,Ni}$ and $P_{cs\,Uj}$ are the rates of sweep-induced coalescent events induced by the *i*th NS site and *j*th UTR site, respectively. The summations are taken over all the sites in the gene that are under selection. The notation S^{-1} is used to denote the reciprocal of the expected time to coalescence due to sweeps, *S*.

Using Equation 4, this expression can be written as:

$$S^{-1} \approx \nu_a \sum_i (B_2\gamma_a)^{-4r_i/s_a} \exp(-2B_1\lambda r_i/s_a)[1+\tfrac{2}{3}(B_1\lambda r_i/s_a)^2]$$
$$+\nu_u \sum_j (B_2\gamma_u)^{-4r_j/s_u} \exp(-2B_1\lambda r_j/s_u)[1+\tfrac{2}{3}(B_1\lambda r_j/s_u)^2] \qquad (5b)$$

where subscripts *a* and *u* denote NS and UTR mutations, respectively.

If we assume that the fixation probability of a favorable mutation in the presence of BGS is discounted by a factor of B_2 compared with the standard value (see above), we have:

$$\nu_a = uB_2p_a\gamma_a \qquad (6a)$$
$$\nu_u = uB_2p_u\gamma_u \qquad (6b)$$

where *u* is the mutation rate per nucleotide site, and p_a and p_u are the proportions of all new NS and UTR mutations, respectively, that are selectively favored.

As described by Campos *et al.* (2017), the summation formula used in the sweep calculations assumes that every third basepair in an exon is a neutral site, with

the other two being subject to selection. This differs from the SLiM procedure of randomly assigning selection status to exonic sites, with a probability p_s of being under selection ($p_s = 0.7$ in the simulations used here). To correct for this, the overall rate of NS substitutions in Equations 5 was adjusted by multiplying by 0.7 x 1.5.

Since we are confining ourselves to a single gene, it is reasonable to assume a linear genetic map. The crossing over contribution to r_i is then given by $r_c z_i$, where z_i is the physical distance between the neutral and selected sites. There is also a contribution from gene conversion, of the same form as for the model of BGS described above.

Following Wiehe and Stephan (1993) and Kim and Stephan (2000), coalescent events caused by selective sweeps and coalescent events caused by neutral drift can be considered as competing exponential processes with rates B_1^{-1} and S^{-1}, respectively. Under the infinite sites model (Kimura 1971), the ratio of expected nucleotide site diversities at a neutral site, relative to the value in the absence of selection at linked sites ($\theta = 4N_e u$, where u is the neutral mutation rate per basepair), can then be written as:

$$\frac{\pi}{\theta} = \frac{1}{B_1^{-1} + S^{-1}} \qquad (7)$$

Partial recovery from sweeps

Equation 7 implicitly assumes a full recovery of diversity within a gene from a sweep, before the next sweep occurs. This assumption is, however, violated if sweeps are sufficiently frequent, and can be relaxed as follows. We assume that sweeps occur in a gene at a constant rate ω per unit of coalescent time, given by the sum over the rates per site for each type of site in the gene. This quantity can be found from Equations 6 by multiplying the rates per site by the number of sites in question. Let the expected neutral diversity at a neutral site immediately after a sweep be π_0, and the expected neutral diversity at the time that the gene experiences a new sweep be π_1. We have:

$$\pi_0 = (1 - D)\pi_1 \qquad (8)$$

where D is the probability that each member of a pair of swept lineages has failed to recombine during the sweep, conditioned on the completion of a sweep. Because the expected reduction in neutral diversity due to recurrent sweeps is S^{-1}, given by Equation 5b, we have $D = (\omega S)^{-1}$, thereby establishing the relationship between π_0 and π_1 (the assumption that the coalescent time for the pair of swept lineages is zero is relaxed below).

Under the infinite sites model ($\theta << 1$), the equilibrium diversity that would be reached in the absence of further sweeps is $B_1\theta$. In this case, the standard formula for the rate of approach of neutral diversity to its equilibrium value (Malécot 1969, p.40), gives the following expression for the diversity at time T after the last sweep:

$$1-\pi(T)(B_1\theta)^{-1} \approx [1-\pi_0(B_1\theta)^{-1}]\exp(-B_1^{-1}T) \qquad (9)$$

(The factor of B_1^{-1} in the exponent reflects the reduction in N_e caused by BGS, resulting in a corresponding acceleration in the rate of approach to equilibrium.)

If sweeps occur at constant rate ω, the probability density of the time of occurrence of the next sweep, T, is given by the exponential distribution, $\omega \exp(-\omega T)$. The expected diversity at the time of the next sweep, π_1, is then given by:

$$\begin{aligned} 1-\pi_1(B_1\theta)^{-1} &\approx [1-\pi_0(B_1\theta)^{-1}]\omega\int_0^\infty \exp[-(\omega+B_1^{-1})T]\,\mathrm{d}T \\ &= [1-\pi_0(B_1\theta)^{-1}]A \end{aligned} \qquad (10)$$

where $A = \omega/(\omega + B_1^{-1})$.

Similarly, the expected diversity over the entire period between successive sweeps, π, is given by:

$$\begin{aligned} 1-\pi(B_1\theta)^{-1} &= [1-\pi_0(B_1\theta)^{-1}]\omega\int_0^\infty \exp(-\omega T)\{T^{-1}\int_0^T \exp(-B_1^{-1}\tau)\,\mathrm{d}\tau\}\,\mathrm{d}T \\ &= [1-\pi_0(B_1\theta)^{-1}]B_1\omega\int_0^\infty \exp(-\omega T)T^{-1}[1-\exp(-B_1^{-1}T)]\,\mathrm{d}T \\ &= [1-\pi_0(B_1\theta)^{-1}]B_1\omega\, I(\omega, B_1) \end{aligned} \qquad (11)$$

Formulae for $I(\omega, B_1)$ are derived in File S1, section 5.

In the absence of any recovery of diversity during the sweep itself, Equations 8-10 together yield the final expression:

$$\frac{\pi}{\theta} = \frac{B_1[1 - B_1\omega ID - A(1-D)]}{1 - A(1-D)} \qquad (12a)$$

In the limit as ω approaches zero, ωI and A both tend to 0, and AD tends to $B_1 S^{-1}$. The value of π/θ for small ω is thus approximately $1/(B_1 + S^{-1})$, corresponding to Equation 7.

To allow for a non-zero mean time to coalescence during the sweep, T_{cs}, the post-sweep diversity π_0 is modified by adding $DT_{cs}\theta$ to Equation 8, where T_{cs} is given by Equation S10. This adds a small additional component to Equation 12a, giving:

$$\frac{\pi}{\theta} = \frac{B_1[1 + \omega ID(T_{cs} - B_1) - A(1-D)]}{1 - A(1-D)} \qquad (12b)$$

Continuum approximation for effects of recurrent sweeps

A useful approximation can be obtained by treating a gene as a continuum, following the treatment of BGS in Campos *et al.* (2017). We correct for the effect of introns simply by reducing the density of NS sites in the coding sequence, by multiplying the density within exons by the fraction of the sites that are exons among the sum of the lengths of the exons, introns and UTRs. In addition, we approximate the effect of gene conversion by writing the net recombination rate between sites separated by z baspairs as $(r_c + g_c)z$ when $z \leq d_g$, and as $r_c z + g$ (where $g = g_c d_g$) when $z > d_g$. The resulting expressions for sweep effects are derived in File S1, section S6. These do not include any corrections for multiple recombination events or for the variance in first passage time, since these make the integrations analytically intractable.

Simulation Results

Effects of background selection alone

Table 2 shows simulation results using the gene model described in the Material and Methods, for chromosomal regions with varying numbers of autosomal loci and varying rates of crossing over, with and without gene conversion. As mentioned in

the Material and Methods, the same rate of gene conversion applies to all the cases with gene conversion, regardless of the rate of crossing over. The estimates of B_1, the ratio of the mean synonymous site nucleotide diversity to the value without selection (θ), are shown in the table, together with their 95% confidence intervals (CIs) over replicate simulations. The mean value of θ from completely neutral simulations was 0.0223, with 95% CI (0.0227, 0.0229), which is slightly lower than the theoretical value on the infinite sites model (0.0239). This probably reflects the fact that SLiM does not distinguish between synonymous and nonsynonymous sites, so there is a probability of 70% that a new mutation at a site segregating for a neutral synonymous mutation will subject to selection, thereby reducing diversity. The mean θ value from the simulations was used to estimate B_1. Table S1 of File S1 shows comparable results for the model of X-linked loci summarized in Table 1, with intermediate dominance and a mean scaled selection coefficient against homozygous deleterious mutations and shape parameter equal to the autosomal values.

Tables 2 and S1 also show the predicted values of B_1 using the continuum model of BGS with the Haldane mapping function described above, using the formulae in File S1, section S1. Equation S3 was numerically integrated over the gamma distribution of selection coefficients, truncated at $\gamma_c = 5$ (see the Material and Methods). The theoretical predictions for the X-linked case are equivalent to those for a mutation rate of ¾ times the autosomal values, with the same values as the autosomal case for all other parameters. Overall, there is a fairly good fit between the theoretical predictions and the simulation results, although the theoretical values of B_1 are mostly smaller than the simulation values.

However, if the additional term in E contributed from neutral mutations that arise in repulsion from a linked deleterious mutation (Equations S1b, S5d and S5e) is ignored, the fits are much less good, especially for the higher rates of crossing over and larger numbers of genes. For example, with 70 autosomal genes and the standard rate of gene conversion, the predicted values of B_1 are then 0.681, 0.790, 0.835, 0.860 and 0.875 for crossover rate factors of 0.5, 1, 1.5, 2 and 2.5, respectively. With 210 genes, the corresponding B_1 values are 0.583, 0.697, 0.739, 0.762 and 0.776; the last value is 20% larger than when the additional term is included.

Similarly, use of a linear relation between physical distance and map distance, which has been assumed in most theoretical models of BGS, generally gives a poorer

fit to the results for the higher rates of crossing over (Table S2 of File S1), except when the number of genes and the map length of the region are both small, reflecting the effect of double crossing over in reducing the net rate of recombination between distant sites. Nonetheless, the fit is surprisingly good overall; indeed, the linear map predictions using Equations S2c, S2d, S4, S5d and S5e often provide a better fit to the simulation results for the cases with 20 and 70 genes. The implications of these effects of the inclusion of the repulsion deleterious mutations, and the difference between the linear and Haldane maps, are considered in the Discussion.

Effects of background selection on the rate of fixation of favorable mutations

The main goal of our work is to analyse the joint effects on neutral diversity of BGS and SSWs, and the extent to which these can be predicted by the relatively simple Equations 7 and 12. A core assumption behind these equations is that the fixation probability of a new favorable mutation is affected by BGS as though N_e is multiplied by a factor that is equal or close to the value that applies to neutral diversity (Kim and Stephan 2000).

We have tested this assumption by comparing the mean numbers of fixations of favorable mutations observed over the last 15,000 generations of the simulations, both without BGS and with BGS. The ratio of these means provides a measure of B (B_2) that can be compared to the value of B estimated from neutral diversity (B_1). There are two reasons why we would not expect perfect agreement. First, a sufficiently strongly selected favorable variant could resist elimination due to its association with deleterious mutations, and instead might drag one or more of them to high frequencies or fixation (Johnson and Barton 2002; Hartfield and Otto 2011). Second, the incursion of selectively favorable mutations may perturb linked deleterious mutations away from their equilibrium, even if they do not cause their fixation.

These Hill-Robertson interference effects (Hill and Robertson 1966; Felsenstein 1974) reduce the N_e experienced by the deleterious mutations, and hence their nucleotide site diversity, which is correlated with the mean number of segregating deleterious mutations. This reduction in the number of segregating deleterious mutations will reduce the effects of BGS on incoming favorable mutations. For both these reasons, B_1 is likely to be smaller than B_2. Table S3 of File S1 provides evidence that the mean number of segregating deleterious mutations is

indeed reduced by selective sweeps, except for the cases with no crossing over, for which the rate of sweeps is greatly reduced compared with cases with crossing over.

This is what is seen in most of the results for autosomal loci shown in Table 4 (Table S4 of File S1 presents some parallel results for X-linked loci). The most extreme case is when there is no crossing over, a regime in which the efficacy of BGS is undermined by Hill-Robertson interference among the deleterious mutations, so that the assumptions underlying the BGS equations tested in the previous section do not apply (McVean and Charlesworth 2000; Comeron and Kreitman 2002; Kaiser and Charlesworth 2009; Seger *et al.* 2010; Good *et al.* 2014; Hough *et al.* 2017). For example, B_1 for 70 genes is 0.086, close to the value found by Kaiser and Charlesworth (2009) for a similar sized region, whereas the standard BGS prediction is 0.04%. In contrast, the B_2 values for favorable NS and UTR mutations are 0.26 and 0.28, respectively, approximately three times greater. This still represents a massive reduction in the efficacy of selection on favorable mutations, consistent with the evidence that their rates of substitution in non-crossover regions of the Drosophila genome are much lower than elsewhere, reviewed by Charlesworth and Campos (2014).

For the other rates of crossing over, there is much closer agreement between the two estimates of B, although the value for favorable mutations is nearly always larger than for neutral mutations. The discrepancy is largest for crossover rates of one-half the standard value, and seems to level off after the standard rate. As might be expected, it is smaller in the presence of gene conversion. It is interesting to note that, in the absence of BGS, there is little effect of the crossing over rate or the number of loci on the rate of fixation of favorable mutations, except for the case of no crossing over, when the rate is substantially lower than for the next lowest rate of crossing over. The extent of this difference increases slightly with the number of genes in the region, with B_1/B_2 for the standard rate of crossing over being 1.02, 1.04 and 1.04 for 70, 140 and 210 genes for NS sites, and 0.95, 1.04 and 1.05 for UTR sites.This suggests that there is interference among selectively favorable mutations when the rate of recombination is very low, but that a relatively low rate of crossing over (of the order of one-half the standard rate) mitigates this effect.

This conclusion is consistent with theoretical predictions for the effects of interfence among positively selected mutations (Kim and Stephan 2003; Weissman and Barton 2012). Equation 4 of Weissman and Barton (2012) gives an approximation

for the rate of substitution of favorable mutations in genomic regions with a linear genetic map of length R, such that ratio of the realized rate of substitution per generation across the region, Λ, to the rate in the absence of interference (Λ_0) is equal to $1/[1 + 2(\Lambda_0/R)]$. If we combine favorable substitutions in NS and UTR sites, which have similar selection coefficients (see Table 1), we have $\Lambda_0 = 0.0249$; for relative crossing over rates of 0.5, and 2.5, $\Lambda /\Lambda_0 = 0.942$ and 0.988, respectively; the ratio of these two values is 0.953. For simulations with no gene conversion, the observed ratio was 0.965, compared with a value of 0.968 with gene conversion. The theoretical expectation of only a small difference between relative rates of crossing over of 0.5 and 2.5 is thus consistent with the simulations, although the quantitative effect is overpredicted by the theory, as was also found by Weissman and Barton (2012, Figure 4). Gene conversion seems to have little effect in reducing the effects of interference in the presence of crossing over.

This formula breaks down in the absence of recombination. Instead, we can use the approximation of Equation 4 of Neher (2013) for the case $s >> U_b$, which is based on Equation 39 of Desai and Fisher (2007). When this is adapted for the case of diploids with semidominance, we have $\Lambda = 0.5s \ln(Ns)/[\ln(2U_b/s)]^2$, where s is the homozygous selection coefficient for a favourable mutation, and U_b is the net favorable mutation rate for the region. Again combining NS and UTR mutations, and putting $s = 0.05$, $U_b = 0.00436$ and $N = 2500$, we have $\Lambda = 0.00406$, and $\Lambda /\Lambda_0 = 0.163$. This can be compared with the observed ratio of the rates of substitution for relative rates of crossing over of 0 and 2.5, with 70 genes and no gene conversion or BGS. This is equal to 0.235, again suggesting that the effect of interference is overpredicted by the approximation. In this case, gene conversion increases the observed ratio to 0.570 (see Table 3), so that it has a considerable effect in mitigating interference when crossing over is absent. BGS seems to play the major role in reducing the rate of substitution of favorable mutations when crossing over is absent, as suggested by Campos *et al.* (2014). The properties of genomic regions with very low rates of crossing over will be analysed in more detail in a later publication.

Effects of selective sweeps on neutral diversity

This section is concerned with four main questions. First, to what extent does partial recovery from sweeps affects the predictions of the models of recurrent sweeps?

Second, how well does the integral approximation for SSWs perform (Equations S24-S33), as compared with the more exact summation formulae (Equations 5 and 6). Third, how well do the competing coalescent process approximations for the joint effects of BGS and SSWs perform, when the various corrections have been included? Finally, is less accuracy obtained by using the neutral BGS value (B_1) rather than B_2 in the appropriate formulae?

Figure 3 shows the relation between γ for favorable mutations, and the predicted ratio of π with the correction for partial recovery of diversity (Equation 11) to its value without this correction (Equation 7). This uses the integral approximation for the effects of selective sweeps. Only NS sites with the mutational parameter values described in the Material and Methods were considered, and the standard rate of crossing over was assumed. No corrections for recovery of diversity or multiple recombination events during a sweep were applied, so that this procedure indicates the extent to which partial recovery affects the accuracy of Equation 7. As might be expected, the effect is strongly dependent on γ. For $\gamma = 250$, the value for NS sites used in the simulations, in the absence of gene conversion there is an approximately 12% reduction in expected diversity when the correction is applied, and an 8% reduction in the presence of gene conversion, with the standard values of the other parameters. BGS slightly reduces the effect of the correction. As would be expected intuitively, the effects increase with γ, in a nearly linear fashion.

Table 4 presents the results of simulations with 70 autosomal genes, together with the predictions for the integral and summation formulae, with and without the corrections. In the case of the corrected summation formulae, all the corrections described above were applied; for the integral results, only the corrections for partial recovery from sweeps and for recovery during sweeps were used. Parallel results for X-linked genes are shown in Tables S5a and S5b of File S1. These involve stronger selection on the favorable mutations, as described in the Material and Methods.

Two types of predictions with BGS are shown. The first applies the B_1 values shown in Table 2 to all the relevant parameters; the second applies B_2 to fixation probabilities, and B_1 to the other parameters. The agreement between the integral and summation results is surprisingly good overall. The largest discrepancies occur when the rate of crossing over is low and there is no BGS, when they are of the order of 4%

of the lower value. This suggests that there should be no substantial loss of accuracy in using the integral approximations in future analyses of selective sweeps.

Table 4 also allows assessment of how well the summation and integral formulae predict the simulation results, and the importance of the corrections for the accuracy of the predictions. For low rates of crossing over and no gene conversion, there are marked discrepancies between the predictions and the simulations when no corrections are applied, especially when there is no gene conversion. For example, in the absence of BGS and gene conversion, the summation formulae predict a value of π/θ for the autosomal case that is 16% greater than the simulation value with a crossing over rate of half the standard value; with BGS and the B_2 correction, the corresponding error is 9.5%. With the standard rate of gene conversion, the errors are 7.4% and 4.3%, respectively. Even with the highest rate of crossing over and gene conversion, which might be expected to mitigate the effects of recurrent sweeps, the values are 2.4% and 4.1% with and without gene conversion, respectively.

With the X-linked results shown in Tables S5a and S5b, which involve more frequent selective sweeps and lower ratios of recombination rates to selection coefficients than in the autosomal case, the discrepancies arising from ignoring partial recovery are even larger. For example, with a rate of crossing over of half the standard value and gene conversion, BGS and the B_2 correction, the error is 13% for the case of the large of the two selection coefficients modeled. With the highest rate of crossing over, it is 7%. These errors arise from the fact that the rates of sweeps per gene are sufficiently high that full recovery between sweeps is unlikely. For the autosomal model, the net rate of sweeps per gene is 1.78 per unit coalescent time in the absence of BGS; the corresponding value for the X-linked model with the larger selection coefficient is 2.67.

It can be seen from the tables that considerable improvements in fit are obtained by applying the corrections. The fit is not, however, perfect, especially for the lowest rate of crossing over with no gene conversion and with BGS. In this case, it seems that using only B_1 overestimates π/θ, whereas the use of B_2 as well as B_1 underestimates it, especially for the integral formulae. The main contribution to the improvements in fit comes from including the effects of partial recovery from sweeps, as can be seen from results where one or both of the other factors (multiple recombination events and recovery of diversity during sweeps) are omitted (Table S6

of file S1). For example, in the absence of BGS but with the standard rate of crossing and neither additional correction, the summation formulae with corrections for partial recovery from sweeps predict π/θ values that are 0.9% and 0.7% greater than the fully corrected values, in the absence and presence of gene conversion, respectively.

With no correction for a recovery of diversity during the sweep, but with a correction for multiple recombination events, the corresponding values are 0.5% less than the fully corrected values. With no correction for multiple recombination events, but with a correction for recovery of diversity during sweeps, the values are 1.5% and 1.1% higher than with the full corrections. Multiple recombination events (which enhance the effects of sweeps) thus have larger proportional effects than the recovery of diversity during sweeps (which reduces sweep effects), but both have smaller effects than the correction for failure to recover after a sweep, at least with the parameters used here. Similarly, the predictions from the integral approximations are only slightly affected by omitting the correction for accrual of diversity during sweeps.

With the standard or higher rates of crossing over, the fits are remarkably good (errors of 2% or less), even with the integral approximation using B_1 for the adjustments to fixation probabilities. Overall, it seems that relatively little is gained by using B_2; for the cases with gene conversion, it gives a worse fit than when only B_1 is used. The values of B_1 from the simulations were used for this purpose. As shown in Table 2, these tend to be somewhat higher than the theoretical predictions described in the first part of the Appendix, but their use does not materially affect the results. For example, with 70 genes and gene conversion, the integral approximation corrected for partial recovery predicts relative diversities of 0.543, 0.626, 0.682, 0.717 and 0.741 for relative rates of crossing over of 0.5, 1, 1.5, 2 and 2.5, respectively. The corresponding simulation results in Table 4 are 0.544, 0.648, 0.703, 0.724 and 0.753. The maximum relative error among these results is 3.4% (for the case relative crossing over rate 1.5); this is somewhat worse than the results obtained in Table 4 using the simulation values of B_1, together with the corrected integral approximation, but is probably acceptable for inference purposes. A heuristic approach is to adjust the theoretical B_1 values upwards by 3%, which is the approximate mean of the relative errors in the theoretical values of B_1 for the case of 70 genes with gene conversion

This yields predicted relative diversity values of 0.544, 0.635, 0.697, 0.735 and 0.760, which are mostly closer to the simulation values.

The predictions of the effects of selective sweeps use a single gene model, based on the assumption that the effects of sweeps with the parameters assumed here are localized to single gene regions. Examination of the simulation results with sweeps alone, displayed in the Supplementary Information (File S2), show that there is no noticeable effect of the numbers of genes on the mean synonymous site diversities, consistent with this assumption. This is not surprising, given that the expected reduction in diversity at a neutral site due to a single sweep at recombination distance r is approximately $\gamma^{-4r/s}$, where γ and s are the scaled and absolute selection coefficients for the favorable allele (Barton 2000). With the values of γ and s for autosomal NS mutations assumed here (250 and 1×10^{-4} for natural populations, respectively), an effective crossing over rate of 1×10^{-8} and a distance of 2000bp between sites (the minimum for sites in separate genes), the expected reduction in diversity in the absence of gene conversion is $250^{(-0.8)} = 0.01$, which is essentially trivial.

This conclusion does not apply when no crossing over is allowed; this case has been studied theoretically by Kim and Stephan (2003) and Weissman and Hallatschek (2014). In this case, the simulation results displayed in File S2 show that there is a large effect of the number of genes. With no crossing over, gene conversion or BGS, the mean autosomal diversities relative to neutral expectation were 0.0819, 0.0700 and 0.0675 for 70, 140 and 210 genes, respectively. These results can be compared to the predictions from the approximate Equation 5 of Weissman and Hallatschek (2014), modified for diploidy with semi-dominance, which gives the absolute neutral nucleotide diversity under recurrent sweeps with recurrent sweeps as $8\mu \ln[2\ln(\gamma)/U_b]/s$. The resulting predicted values are 0.195, 0.183 and 0.176, respectively. As was also found by Weissman and Hallatschek (2014), these considerably overpredict the diversities. Gene conversion greatly reduces the effects of sweeps, with relative diversities of 0.130, 0.090 and 0.0832. However, BGS has a much greater effect on diversity than sweeps; with gene conversion, it gives relative diversity values of 0.0867, 0.0429 and 0.0293 for 70, 140 and 210 genes, respectively. Essentially the same values are seen with both BGS and SSWs, reflecting the fact that the rate of sweeps is greatly reduced in the presence of BGS (see Table 3). The

predicted relative diversity value for a 70 gene region is quite close that observed for the fourth chromosome of *D. melanogaster*, which has a comparable number of genes (Campos *et al.* 2014), suggesting that diversity in non-crossover regions of the genome is largely controlled by BGS, as was also inferred by Hough *et al.* (2017) for the case of the newly evolved Y chromosome of *Rumex*.

Discussion

Accuracy of the approximations for pairwise diversity with hitchhiking

We have developed a new approximation for the effect of a single selective sweep on pairwise neutral diversity at a site linked to the target of selection (Equation 4). This uses an approximate formula for the duration of a sweep, which includes the contribution from the stochastic phases of the sweep (Equations 2 and 3). In addition, we have developed corrections for the accrual of diversity at a neutral site that remains associated with the selectively favorable allele during the course of the sweep, and for multiple recombination events during the sweep (sections S3 and S4 of File S1). While these problems have been studied previously by more exact methods, e.g. Stephan *et al.* (1992), Barton (1998, 2000) and Hermisson and Pennings (2005), the simulation results in Tables 4 and S5 suggest that the heuristic approach used here provides adequate approximations.

More importantly, we have derived formulae that allow predictions of the effects of recurrent sweeps on pairwise neutral diversity (Equations 12), on the assumption that these occur at a constant expected rate. This relaxes the assumption made in most previous models of recurrent sweeps that diversity is fully recovered after one sweep before a gene is hit by the next sweep, based on Equation 11 (Wiehe and Stephan 1993; Kim and Stephan 2000; Coop and Ralph 2012)}. This expression has been used several times in making inferences about sweep parameters from population genomic data (Sella *et al.* 2009; Elyashiv *et al.* 2016; Campos *et al.* 2017), so that improved accuracy from such inferences should result by using the more general expressions derived here. As described in the previous section, our simulation results, which use parameter values that are consistent with those estimated for *D. melanogaster* by Campos *et al.* (2017), suggest that the use of Equations 12 considerably improves the fit of the predictions based on the theoretical formulae,

reflecting the fact that sweeps are sufficiently frequent that full recovery of diversity between sweeps cannot occur. While the results of Kim *et al.* (2006) also relax the assumption of no recovery between sweeps, and provide expressions for the site frequency spectrum at neutral sites affected by sweeps, these involve complex calculations, in contrast to our fairly simple expressions.

The comparisons of the simulation results with the theoretical predictions also suggest that the correction for accrual of diversity during the sweep itself has a relatively small effect (Table S6), so that the simpler Equation 12a that ignores the correction for diversity recovery is adequate for most practical purposes. Similarly, while the correction for multiple recombination events between the selectively favorable allele and a neutral site (Equations S21) has a somewhat larger effect, is nevertheless sufficiently small that it can probably be ignored.

While our simulations have involved 'hard' sweeps, where the new favorable mutation is introduced as a single copy, Equation 12a can also be applied to other situations, such as 'soft' sweeps arising from standing variation or multiple mutations to the favorable allele at a locus (Hermisson and Pennings 2005, 2017). The only modification that need be made is to the expression for the reduction in diversity immediately after a sweep (D) in Equation 8.

Another feature of the work presented here is that we have incorporated gene conversion into sweep models, as was also done by Campos *et al.* (2017), but which has been ignored in previous treatments of sweeps. Gene conversion events that are not associated with crossovers are known to be a major source of recombination events at the intragenic level in *Drosophila* (Hilliker and Chovnick 1981; Hughes *et al.* 2018). With the standard autosomal effective crossing over rate for *D. melanogaster* of 1×10^{-8} per bp (Campos *et al.* 2014), the effective rate of crossing over between two sites separated by 500bp is 5×10^{-6}. With scaled and absolute selection coefficients for NS mutations of $\gamma = 250$ and $s = 10^{-4}$, as assumed in the discussion of Table 4 at the end of the Results section, and in the absence of gene conversion, the expected proportional reduction in diversity at the end of a sweep for a neutral site that is 500bp away from the selected site is approximately $\gamma^{(-4r/s)} = 250^{(-4 \times 0.05)} = 0.33$. With the gene conversion parameters assumed here, which are based on experimental estimates from *D. melanogaster* (Hilliker and Chovnick 1981; Hughes *et al.* 2018), use of the formula of Frisse *et al.* (2001) shows there is a

additional contribution to the net effective rate of recombination of 3×10^{-6}, so that the total effective recombination rate is 8×10^{-6}. This yields a reduction in diversity of 0.17, approximately 50% of the value in the absence of gene conversion. Consistent with this result, the simulation results and theoretical predictions are significantly affected by gene conversion, such that the expected effects of sweeps on diversity are considerably reduced if gene conversion is present (Tables 4 and S5). Ignoring gene conversion is thus likely to substantially bias estimates of sweep parameters.

It is also of interest to note that, as noted at the end of the Results section, the use of these approximations, together with the reduction in diversity at neutral sites caused by BGS (the predictions using B_1, described in the Material and Methods), provide quite adequate predictions. The results suggest that inference methods can be simplified by using the integral approximations to both selective sweeps and BGS, rather than the summation formulae used in Equations 5.

The relation between sequence diversity and rate of crossing over

It is also of interest to ask what light the theoretical results described above shed on the observed positive relationship between DNA sequence variability at putatively neutral or nearly neutral sites within a gene in *D. melanogaster* and the local rate of recombination experienced by the gene (Aguadé *et al.* 1989; Begun and Aquadro 1992). This observation stimulated interest in models of SSWs and BGS, and its cause has been a long-standing subject of debate; for reviews, see Sella *et al.* (2009), Stephan (2010), Cutter and Payseur (2013) and Charlesworth and Campos (2014). We first note that recent analyses of population genomic data suggest a strong relationship between synonymous nucleotide site diversity (π_S) and the effective rate of crossing over, provided that genes in regions with no crossing over are excluded. For example, Figure S2 of Campos *et al.* (214) presents estimates of population genetic parameters for a sample from a Rwandan population of haploid genomes of *D. melanogaster* for bins of genes as functions of the mean effective rate of crossing over for each bin, obtained from the data of Comeron *et al.* (2012), shows that the mean autosomal π_S for a sample from a Rwandan population of *D. melanogaster* increases from approximately 0.0083 to 0.0192 as the effective rate of crossing increases from 0.5 cM/Mb to 2 cM/Mb, which is the upper limit to the estimate autosomal rate of crossing over (these rates correspond to the relative rates of crossing over of 0.5 and 2

used above). The ratio of the two values is 2.31. The simulation results for both BGS and SSWs with 70 genes and gene conversion, shown in Table 4, give a ratio of 1.33; with BGS alone, the ratio is 1.20, and with SSWs alone it is 1.24 (the ratios are only slightly affected by the number of genes in the region, as can be seen from the results in File S2).

This raises the question of what causes this discrepancy. One possibility is that the mean scaled selection coefficients for favorable mutations used in these simulations are unrealistic. This was checked by re-running the calculations with different γ values for the favorable mutations, using the integral approximation with the correction for partial recovery, the standard rate of gene conversion with SSWs and BGS, and using only B_1 for all relevant BGS parameters. This gives a ratio of 1.29 for the standard γ values; the ratios for γ values that are half and twice these, are 1.28 and 1.41, respectively. There is thus only a weak dependence on the strength of selection on favorable mutations. This is not surprising, in view of the fact that the natural logarithm of the net effect of sweeps on a neutral site for favorable mutations, with scaled selection coefficient γ and scaled recombination rate ρ, is approximately equal to $\ln(\gamma)[1 - 4\rho\gamma^{-1}]$ plus a constant. For $\gamma >> 1$, the derivative of this expression with respect to γ is approximately equal to $\gamma^{-1} [1 + 4\rho\gamma^{-1} \ln(\gamma)]$, which means that there is only a small proportional effect on diversity of a change in γ, for a given value of ρ. Similarly, its derivative with respect to ρ to is $- 4\gamma^{-1} \ln(\gamma)$, which is $<< 1$ when $\gamma >> 1$. It thus seems unlikely that the weak dependence of neutral diversity on the rate of crossing over can be explained by the choice of selection coefficients for favorable mutations.

The effect of the proportion of mutations that are beneficial can be examined in a similar way. Halving these leads to a ratio of 1.28 for the diversities at relative crossing over rates of 2 and 0.5, and doubling them to a ratio of 1.41. Although the parameter has a large effect on the absolute diversity levels, its proportional effects on the values for different relative crossing over rates are nearly independent of the crossing over rates. Even if both the strengths of selection and the proportions of beneficial mutations are doubled, the ratio of diversity values is increased to only 1.69 without BGS. In the absence of BGS, however, the ratio for this case is 1.84, presumably because the absence of BGS allows a faster rate of fixation of favorable mutations. To explain the observed relation between diversity and rate of crossing,

considerably larger values of both the strength of selection and proportion of favorable mutations than are currently suggested by population genomic analyses seem to be required.

Another possibility is that intergenic and intronic sequences are subject to selection, rather than being selectively neutral. Charlesworth (2012b) used evidence on the levels of selective constraints on different types of *Drosophila* DNA sequences to obtain crude estimate of γ values for deleterious mutations in weakly constrained and strongly constrained noncoding sequences, as well as for deleterious NS mutations. His analysis showed that a linear genetic map provided a good approximation to the BGS predictions. We have therefore used this approach to predict the background selection parameter B_1 for a genic region with a given rate of crossing over, modifying it slightly to include the effect of gene conversion. The details are described in File S1, section S7. For a model of an autosome with the standard rate of gene conversion, this procedure gives B_1 values of 0.340, 0.583, 0.698, 0.763 and 0.806 for relative rates of crossing over of 0.5, 1, 1.5, 2 and 2.5, respectively, yielding a ratio of 2.24 for B_1 for relative crossing over rates of 2 and 0.5. If these values are then applied to the integral approximation for SSWs with the selection parameters used in Table 4 and the standard rate of gene conversion, the predicted values of π/θ for these relative rates of crossing over become 0.305, 0.531, 0.622, 0.679 and 0.725, respectively; this gives a ratio of 2.23 for relative rates of 2 and 0.5.

The same procedure can also be applied to the X chromosome. In this case, we obtain values of B_1 values of 0.507, 0.712, 0.797, 0.844 and 0.873 for relative rates of crossing over of 0.5, 1, 1.5, 2 and 2.5, respectively. The ratio of π values for relative crossover rates of 2 and 0.5 is now only 1.66, which implies a much shallower relationship between π_S and the rate of crossing over than for the autosomes, reflecting the higher effective rate of recombination on the X chromosome. If these value are used to predict the ratio of π values for relative crossover rates of 0.5 and 2 with SSWs, a value of 1.59 is obtained with both the lower and higher selection coefficient for favorable X-linked mutations used previously (see the Material and Methods). Qualitatively at least, this pattern matches the lower slope for plots of X diversity against the effective rate of crossing over; the procedure that was used for the autosome gives an observed ratio of approximately 1.63 for the ratio of π_S values

with effective rates of crossing over of 0.5 and 2cM/Mb. Similarly, assuming that the value of π_S for X-linked mutations in the absence of hitchhiking effects is three-quarters of the value for autosomal mutations, the X/A diversity ratios for relative rates of crossing over of 0.5 and 2 with SSWs and BGS with strong selection on favorable X-linked mutations should be 0.956 and 0.677 with the weaker X-linked selection coefficient, and 0.848 and 0.600 with the stronger X-linked selection coefficient; the observed values are approximately 1.10 and 0.823, respectively. This suggests that the X/A ratio of N_e may in fact be somewhat greater than three-quarters, perhaps reflecting the effect of sexual selection (Charlesworth 2001).

Similar calculations can be performed for estimates of ω_a, the rate of substitution of favourable NS mutations relative to the neutral rate. The expected relative values of ω_a for different rates of recombination should approximately reflect the corresponding relative values of B_1, give the evidence that there is little interference among positively selected mutations. For autosomes, this predicts a ratio of 2.24 for crossing over rates of 2 and 0.5, compared with an observed value of 2.75. For the X chromosome, the predicted relative value is 1.66, and the observed value 1.62. Given the uncertainties in the individual estimates of ω_a, there is reasonably good agreement between the predicted and observed values. Castellano *et al.* (2016) suggested that the smoothing procedure used by Campos *et al.* (2014) to estimate rates of crossing over for each bin might produce biased results, and instead conducted analyses of the relation between the rate of adaptive evolution and unsmoothed rates of crossing over obtained from Comeron *et al.* (2012). For the Rwandan sample used here, with the same number of bins of rates of crossing over (but including non-crossover regions), their non-linear regression equation (line 5 of their Table 1) yields a ratio of 1.95 for the estimated rates of adaptive evolution for relative effective rates of crossing over 2 and 0.5, which is somewhat closer to the above prediction than the value given above. However, the inclusion of the non-crossover genes by Castellano *et al.* (2016) may exaggerate the curvilinearity of the regression equation as applied to crossover regions, given the special properties of non-crossover regions.

These analyses are obviously quite crude, but suggests that the relative values of nearly neutral variability in crossover regions of the *D. melanogaster* genome mainly reflect the effects of BGS rather than SSWs, in agreement with Comeron

(2014). As discussed at the end of the Results section, this conclusion also applies to regions of the genome with zero or very low rates of crossing over, where the effects of SSWs are expected to be weak.

Distortion of the site frequency spectrum by hitchhiking

We have not previously discussed the effects of BGS and SSWs on the site frequency spectra (SFS) at the neutral loci affected by selection at linked sites in genomic regions with crossing over. While it should be possible to use the theoretical frameworks developed for BGS (Zeng and Charlesworth 2011; Nicolaisen and Desai 2013) and SSWs (Durrett and Schweinsberg 2004; Kim 2006 ; Pfaffelhuber *et al.* 2006; Bossert and Pfaffelhuber 2013), this would require extensive calculations that are outside the scope of this paper. We note, however, that the simulation results shown in File S2 show that recurrent SSWs have quite strong effects on the SFS, even with quite high rates of crossing over, in the direction of an excess of rare variants over neutral expectation. This is expected from previous theoretical work (Kim 2006), and has been seen in previous simulation studies, e.g. Messer and Petrov (2013).

For example, with 70 autosomal genes and gene conversion, the mean values of synonymous site Tajima's *D* with SSWs and BGS for relative rates of crossing over of 0.5, 1, 1.5, 2.0, and 2.5 were – 0.209, – 0.156, – 0.116, – 0.111 and – 0.069, respectively. The corresponding mean proportions of singletons were 0.319, 0.310, 0.302, 0.299 and 0.295, compared with the neutral value from simulations of 0.275. In the presence of BGS but not SSWs, the mean values of Tajima's *D* were – 0.046, – 0.013, – 0.019, – 0.036 and 0.000, respectively, compared with the neutral value of 0.042. The mean values of the proportions of singletons were 0.288, 0.286, 0.284, 0.289 and 0.282. Thus, with the parameters used here, BGS contributes very little to the distortion in the SFS, as is expected from previous theoretical work on BGS with significant amounts of recombination (Zeng and Charlesworth 2011; Nicolaisen and Desai 2013). Detailed comparisons with the data are made difficult by the probable effects of demographic factors on these measures of distortion of the SFS, which will tend to obscure the effects of selection at linked sites, especially their relations with the rate of crossing over.

As might be expected, stronger selection on favorable mutations increases the extent of distortion of the SFS. For example, with the stronger of the two selection

models for the X chromosome, the Tajima's *D* values and proportions of singletons for the standard rate of crossing over for 70 genes with gene conversion, SSWs and BGS were – 0.434 and 0.360, respectively. The difference between X and autosomes is qualitatively similar to what is seen for the Rwandan population of *D. melanogaster*, shown in Figure 4 of Campos *et al.* (2014). However, the distortion of the SFS on the X chromosome is much greater than is ever seen in the simulations. It remains to be seen whether more complex demographic scenarios than the constant population size assumed here can explain this discrepancy.

The picture is, however, very different when crossing over is absent. For 70 autosomal genes with gene conversion, the means of Tajima's *D* and the proportion of singletons for synonymous sites with BGS alone were – 0.880 and 0.488, respectively. With SSWs as well, the values were changed by relatively small amounts, to – 1.306 and 0.563, respectively, reflecting the greatly reduced rate of fixations of favorable mutations when crossing over is absent (Table 3). It therefore seems likely that the distorted SFSs seen in genomic regions that lack crossing over (Cutter and Payseur 2013; Campos *et al.* 2014) are mainly caused by BGS in the weak interference selection limit, when interference among sites subject to purifying selection causes genealogies at linked sites to have longer terminal branches (Gordo *et al.* 2002; Kaiser and Charlesworth 2009; Seger *et al.* 2010; O'Fallon *et al.* 2010; Good *et al.* 2014).

Problems with simulating BGS

We conclude with a discussion of some more technical questions concerning the modelling of BGS in SLiM. As described in the first part of the Results section, the fact that SLiM assumes a lack of crossover interference requires the modification of the standard BGS equations to model the Haldane mapping function; this is described in the first part of the Appendix. In addition, for accurate approximations to the simulation results, it was necessary to include an additional term in the BGS equations that results from deleterious mutations that were in initially in repulsion with a new neutral variant (Santiago and Caballero 1998; Charlesworth 2012b), which is ignored in the equations that are usually used to model BGS (see the discussion of Table 2).

These features of the results are more a reflection of the simulation procedure than of biological reality. Equation S1b implies that the extra term added to the standard BGS equation of Nordborg *et al.* (1996) is proportional to the sum of twice the product of the deleterious mutation rates and the mean of *hs* for deleterious mutations, multiplied by a term that is nearly independent of the factor used for rescaling. This term is exactly equal to this product when there is no recombination, and is then equal to the additive variance for fitness under deterministic mutation-selection balance (Mukai *et al.* 1972). Since the deterministic parameters that are thought to be realistic for a Drosophila population have been multiplied by 532 for use in the simulations, the additive genetic variance in fitness is multiplied by a factor of $(532)^2 = 283{,}024$ compared with its value for the real population. With 70 genes, for example, the additive variance in the simulations is 0.0542, whereas the corresponding value for the population is 1.92×10^{-7}. In contrast, the Nordborg *et al.* (1996) equation depends largely on the ratios of deterministic parameters, except for the multiplication of the recombination rate by a factor of $1 - hs$, and so is largely unaffected by the rescaling. In the real population, this additional term is effectively negligible, justifying the use of the standard equation for modeling BGS, e.g. (McVicker *et al.* 2009; Charlesworth 2012b; Comeron 2014; Elyashiv *et al.* 2016; Campos *et al.* 2017).

The use of the Haldane mapping function also means that the simulated rate of recombination for the region as a whole is affected by the rescaling, since the frequency of double crossovers is greatly increased over what would be found in a region of the same physical length in the real population. For example, with the standard rate of crossing over and 70 genes, the map length of the region with the standard rate of crossing over is 1.62. With a Poisson distribution of numbers of crossovers, as assumed in the simulations, the proportion of double crossovers among chromosomes that have experienced a crossover is $0.5 \times (1.62)^2 \times \exp(-1.62)/[1 - \exp(-1.62)] = 0.324$. For regions of the size that we have simulated, the high level of crossover interference that occurs in *Drosophila* (Hughes *et al.* 2018) means that a linear relation between the frequency of crossing over and physical distance is close to reality for the parameters that apply to the real population (Charlesworth 2012b). Unfortunately, except for the cases with a frequency of crossing over of one-half the standard rate used here, it is impossible to simulate a linear model with 70 genes or more, since the expected number of crossovers in the region is greater than one,

which is inconsistent with a probability model that assumes that there is either a crossover or no crossover in the region.

Given that our simulation results generally support the use of the theoretical formulae for both background selection and selective sweeps, largely because both BGS and SSW effects extend over much smaller distances than the whole region, this implies that the use of formulae based on the BGS and SSW equations with a linear genetic map is probably justified for most analyses of population genomic data, although it would be desirable to validate this conclusion with simulations using much larger population sizes than was feasible here. New approaches to forward simulations suggest that this should shortly become possible without an undue expenditure of computer time (Kelleher *et al.* 2018).

Acknowledgments

This work was supported by grant RPG-2015-2033 from the Leverhulme Trust (to BC). We thank Hannes Becher for useful discussions and comments on the manuscript.

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Table 1 Parameters used in the simulations

	Natural population		Simulations	
Parameter	**A**	**X**	**A**	**X**
Population size (*N*)	1.33×10^{6}	0.997×10^{6}	2500	2500
Rescaling factor	-		532	532
Standard effective crossover rate	1×10^{-8}	1.33×10^{-8}	5.32×10^{-6}	5.32×10^{-6}
G.c. rate of initiation	1×10^{-8}	1.33×10^{-8}	5.32×10^{-6}	5.32×10^{-6}
G.c. tract length	440 bp	440bp	440 bp	440 bp
Mutation rate per bp	4.5×10^{-9}	4.5×10^{-9}	2.39×10^{-6}	1.79×10^{-6}
γ_{NS}	2000	20000	2000	2000
γ_{UT}	110	110	110	110
γ_a	250	375 or 500	250	375 or 500
γ_u	213	319.5 or 416	213	319.5 or 416
p_a	2.21×10^{-4}	2.21×10^{-4}	2.21×10^{-4}	2.21×10^{-4}
p_u	9.04×10^{-4}	9.04×10^{-4}	9.04×10^{-4}	9.04×10^{-4}
Prop. of neutral exonic mutations	0.3	0.3	0.3	0.3
Shape parameter of gamma distribution	0.3	0.3	0.3	0.3
Dominance coefficient	0.5	0.5	0.5	0.5

See text for meaning of the parameters.

Table 2 BGS predictions and simulation results for autosomal values of $B_1 = \pi/\theta$

Xover Rate	**0.5**	**1.0**	**1.5**	**2.0**	**2.5**
No g.c					
20 genes	0.687, 0.723 (0.702, 0.745	0.798, 0.838 (0.823, 0.852)	0.845, 0.861 (0.843, 0.879)	0.871, 0.868 (0.849, 0.887)	0.888, 0.913 (0.894, 0.933)
70 genes	0.592, 0.643 (0.623, 0.645)	0.716, 0.737 (0.721,0.750)	0.767, 0.790 (0.770, 0.799)	0.794, 0.818 (0.810, 0.827)	0.812, 0.830 (0.824, 0.838)
140 genes	0.514, 0.543 (0.534, 0.550)	0.632, 0.655 (0.648, 0.663)	0.679, 0.701 (0.694, 0.709)	0.704, 0.723 (0.717, 0.729)	0.719, 0.731 (0.724, 0.739)
210 genes	0.452, 0.489 (0.481, 0.491)	0.559, 0.582 (0.574, 0.590)	0.601, 0.620 (0.618, 0.624)	0.623, 0.642 (0.638, 0.646)	0.637, 0.654 (0.650, 0.658)
G.c.					
20 genes	0.753, 0.796 (0.775, 0.819)	0.836, 0.883 (0.862, 0.903)	0.872, 0.905 (0.889, 0.920)	0.892, 0.907 (0.887, 0.929)	0.905, 0.924 (0.905, 0.942)
70 genes	0.650, 0.686 (0.677, 0.693)	0.750, 0.782 (0.767, 0.799)	0.791, 0.816 (0.797,0.834)	0.813, 0.820 (0.813,0.827)	0.827, 0.838 (0.830,0.848)
140 genes	0.563, 0.594 (0.588, 0.601)	0.662, 0.687 (0.676, 0.696)	0.700, 0.719 (0.707,0.728)	0.720, 0.725 (0.720,0.731)	0.733, 0.736 (0.729,0.744)
210 genes	0.496, 0.525 (0.519,0.531)	0.586, 0.605 (0.598,0.613)	0.620, 0.640 (0.635,0.645)	0.638, 0.639 (0.635,0.640)	0.649, 0.657 (0.652,0.661)

The left-hand upper entries in the cells show the predicted values of B_1, the ratio of the mean synonymous site diversity with BGS (but no sweeps) to its value in the absence of BGS (using Equations S1, S2a, A2b, S3, and S5), and integrating over the truncated gamma distribution. The right-hand upper entries are the corresponding observed mean values. The lower entries are the lower and upper 2.5 percentiles of the observed values of B_1, obtained from the means of the synonymous site diversities over the entire region for each replicate simulation.

The rows labelled 'Xover rate' refer to the results for rates of crossing over with ratios of 0.5, 1, 1.5, etc. to the standard rate of 5.32×10^{-6} used in the simulations.

Cases with no gene conversion are denoted by 'No g.c.' and cases with the standard gene conversion parameters are labelled 'G.c.'.

Table 3 The effect of BGS on the numbers of fixations of selectively favorable autosomal mutations

Gene No.	Xover Rate	No BGS	With BGS	Ratio (B_2)	B_1
70	0	1.23 (1.17,1.27) 1.58 (1.53,1.62)	0.32 (0.30,0.35) 0.45 (0.42,0.48)	0.263±0.012 0.285±0.017	0.086
	0.5	1.83 (1.77,1.89) 2.94 (2.87,3.02)	1.38 (1.34,1.43) 2.14 (2.08,2.20)	0.754±0.018 0.726±0.014	0.686
	1.0	1.97 (1.87,2.08) 3.10 (2.95,3.26)	1.57 (1.46,1.70) 2.28 (2.13,2.45)	0.797±0.019 0.735±0.021	0.782
	1.5	1.96 (1.88,2.04) 2.94 (2.84,2.99)	1.63 (1.52,1.74) 2.47 (2.26,2.47)	0.831±0.023 0.838±0.021	0.816
	2.0	1.88 (1.82,1.94) 3.04 (2.80,3.04)	1.59 (1.53,1.65) 2.49 (2.47,2.57)	0.844±0.021 0.820±0.018	0.820
	2.5	1.89 (1.83,1.96) 3.04 (2.97,3.11)	1.60 (1.53,1.67) 2.44 (2.37,2.52)	0.845±0.024 0.803±0.015	0.838
140	0	0.90 (0.88,0.93) 1.16 (1.12,1.19)	0.10 (0.09,0.11) 0.13 (0.12,0.14)	0.111±0.006 0.112±0.004	0.043
	0.5	1.87 (1.83,1.92) 2.88 (2.82, 2.99)	1.24 (1.20,1.28) 2.49 (2.42,2.56)	0.659±0.013 0.865±0.017	0.594
	1.0	1.97 (1.90,2.03) 2.95 (2.86,3.05)	1.41 (1.35,1.47) 2.10 (2.02,2.16)	0.717±0.020 0.712±0.017	0.687
	1.5	1.91 (1.85,1.96) 3.01 (2.92,3.09)	1.39 (1.32,1.47) 2.22 (2.15,2.30)	0.728±0.023 0.734±0.017	0.719
	2.0	1.88 (1.84,1.92) 2.95 (2.90,3.00)	1.42 (1.37,1.46) 2.20 (2.15,2.26)	0.752±0.015 0.746±0.012	0.725
	2.5	1.96 (1.92,2.01) 2.98 (2.92,3.05)	1.42 (1.37,1.47) 2.15 (2.10,2.21)	0.723±0.015 0.722±0.012	0.736
210	0	0.75 (0.73,0.77) 0.95 (0.92,0.97)	0.05 (0.04,0.06) 0.07 (0.06,0.08)	0.072±0.007 0.076±0.005	0.029
	0.5	1.86 (1.81,1.90) 2.86 (2.80,2.91)	1.09 (1.06,1.13) 1.66 (1.62,1.70)	0.587±0.012 0.591±0.009	0.525
	1.0	1.87 (1.84,1.91) 2.90 (2.84,2.97)	1.18 (1.10,1.25) 1.85 (1.80,1.91)	0.631±0.021 0.638±0.012	0.605
	1.5	1.85 (1.80,1.90) 2.91 (2.86,2.98)	1.22 (1.17,1.26) 1.89 (2.86,2.98)	0.659±0.015 0.679±0.011	0.640
	2.0	1.89 (1.84,1.93) 2.98 (2.93,3.03)	1.27 (1.25,1.30) 1.95 (1.92,1.99)	0.676±0.008 0.655±0.008	0.639
	2.5	1.92 (1.88,1.96) 2.94 (2.89,2.99)	1.26 (1.23,1.29) 2.01 (1.97,2.05)	0.655±0.010 0.684±0.009	0.657

The upper and lower entries in the cells in the third and fourth columns show the ratios of the mean numbers of fixations (over the final 15,000 generations of the simulations) to the number of simulated genes, for selectively favorable NS and UTR mutations, respectively. The fifth column shows the ratios of these values for simulations with and without BGS, respectively, with approximate standard errors calculated from the upper and lower 2.5 percentiles of the numerator and denominator (the percentiles for the ratios are not given, since the ratios are not normally distributed).

The B_1 values in the last column were obtained from Table 2.

The standard gene conversion parameters are assumed.

Table 4 Observed and predicted values of autosomal neutral diversity for a 70 gene region, relative to the value without hitchhiking effects

Xover Rate	Observed	Integral, NC	Sum., NC	Integral, C	Sum., C
No g.c.					
0.5	0.516 (0.500,0.528) 0.430 (0.419,0.441)	0.585 0.489 0.463	0.601 0.496 0.471	0.495 0.463 0.378	0.515 0.471 0.390
1.0	0.655 (0.637,0.671) 0.555 (0.536,0.573)	0.710 0.596 0.591	0.727 0.604 0.600	0.645 0.547 0.535	0.665 0.559 0.547
1.5	0.735 (0.727,0.743) 0.631 (0.621,0.643)	0.782 0.667 0.662	0.797 0.676 0.671	0.732 0.625 0.618	0.750 0.636 0.630
2.0	0.772 (0.763,0.781) 0.675 (0.666,0.683)	0.827 0.713 0.709	0.840 0.721 0.717	0.786 0.678 0.673	0.803 0.689 0.684
2.5	0.812 (0.812,0.828) 0.715 (0.706,0.724)	0.857 0.741 0.737	0.869 0.748 0.745	0.824 0.712 0.707	0.838 0.722 0.717
G.c.					
0.5	0.685 (0.674,0.695) 0.544 (0.534,0.552)	0.750 0.584 0.577	0.736 0.577 0.569	0.693 0.558 0.529	0.675 0.549 0.519
1.0	0.767 (0.763,0.771) 0.648 (0.626,0.660)	0.817 0.682 0.675	0.811 0.678 0.676	0.774 0.647 0.643	0.767 0.641 0.638
1.5	0.815 (0.809,0.821) 0.703 (0.690,0.717)	0.854 0.728 0.726	0.853 0.727 0.725	0.819 0.699 0.697	0.818 0.697 0.695
2.0	0.850 (0.834,0.856) 0.724 (0.713,0.736)	0.878 0.746 0.744	0.879 0.747 0.745	0.849 0.722 0.720	0.851 0.723 0.721
2.5	0.863 (0.858,0.869) 0.753 (0.744,0.761)	0.895 0.771 0.770	0.898 0.773 0.771	0.870 0.750 0.748	0.874 0.752 0.750

The entries for the observed values are the mean synonymous site diversities from the simulations with 70 genes, measured relative to the corresponding values in the absence of selection at linked sites. The upper and lower entries in each cell are the values with SSWs alone and with SSWs and BGS, respectively.

The upper entries in each cell for the predictions are the reductions with SSWs alone; the middle entries include the BGS effect for neutral sites(B_1); the lowest entries include the BGS effects for NS selected site effects (B_2) where relevant. (B_2 values for NS and UTR sites are similar.) The B_2 values with a crossing over rate of half the standard value were estimated separately; the values for the other crossing over rates were pooled.

The columns labelled ‘Integral’ use the approximate integral formulae for SSW effects (equations S24-33); those labelled ‘Summ.’ use the summation formulae, Equations 5 and 6.
‘NC’ denotes predictions without correcting for partial recovery from sweeps (Equation 5).
‘C’ denotes predictions that correct for partial recovery (Equations 12).
For the summation predictions, corrections for multiple recombination events during the sweep (Equation S21b) and for the variance in first arrival time during the first stochastic phase (Equation S8C) were applied.
‘No g.c.’ and ‘G.c.’ refer to results without gene conversion and with the standard gene conversion parameters, respectively.

FIGURE LEGENDS

Figure 1 The gene model used in the simulations.

Figure 2. The possible fates of pairs of neutral lineages sampled after sweep, with no recombination on the left, and two recombination events on the left.

Figure 3 The relation between the scaled selection coefficient on a favorable mutation (γ) and the ratio of the expected neutral diversity with a correction for partial recovery to its value with no correction (Equation 7). The continuum approximation for sweep effects (Equations S24-S33) was used, with the standard parameters of selection, crossing over and gene conversion described in the Material and Methods. The full lines are for the case with no effect of BGS, and the dashed lines assume a value of $B_1 = B_2 = 0.75$. The scale for γ runs from 0 to 1000.

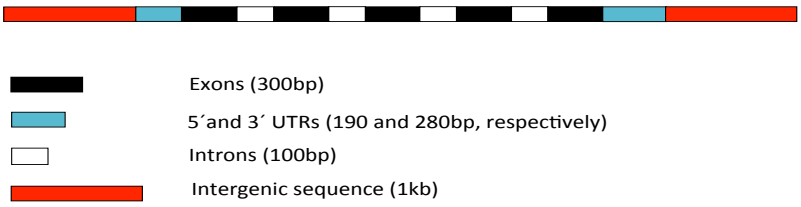

Fig. 1

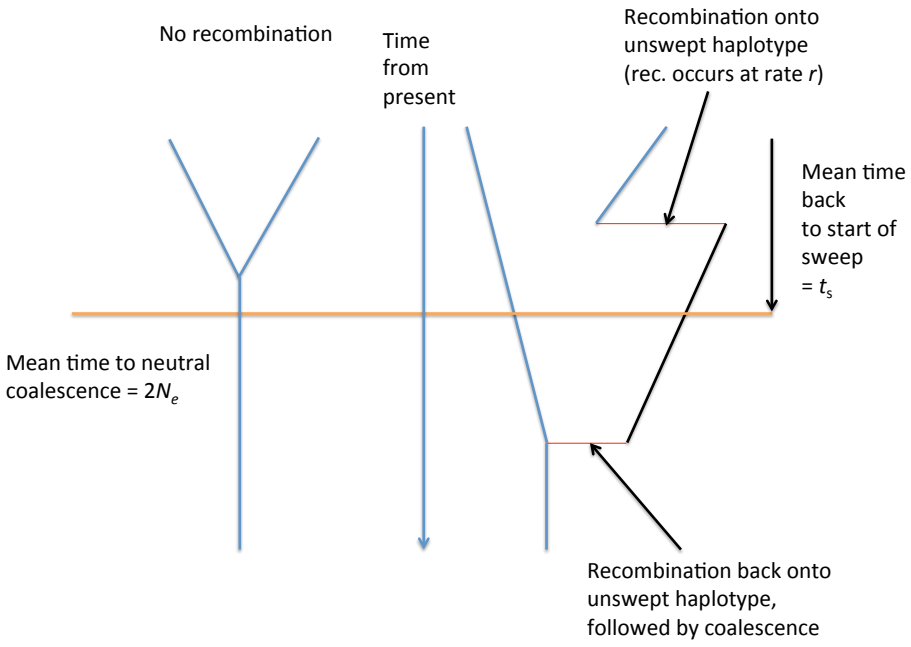

Fig. 2

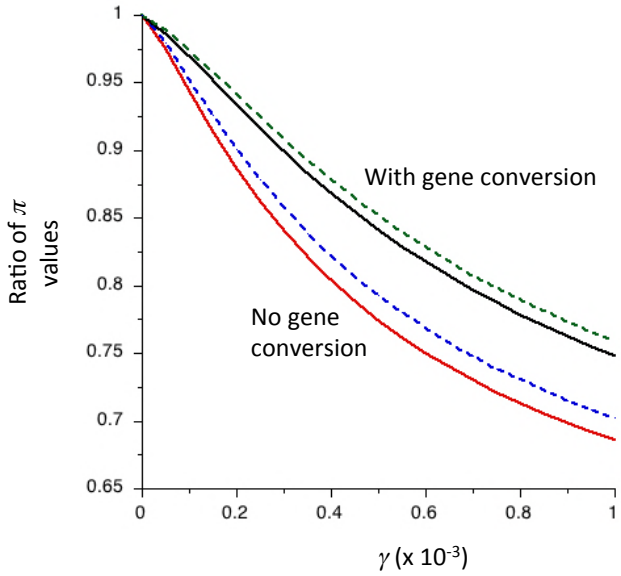

Fig. 3