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Scaling limits for critical inhomogeneous random graphs with finite third moments*

Shankar Bhamidi[†] Remco van der Hofstad[‡] Johan S.H. van Leeuwaarden[‡]

Abstract

We identify the scaling limit for the sizes of the largest components at criticality for inhomogeneous random graphs with weights that have finite third moments. We show that the sizes of the (rescaled) components converge to the excursion lengths of an inhomogeneous Brownian motion, which extends results of Aldous [1] for the critical behavior of Erdős-Rényi random graphs. We rely heavily on martingale convergence techniques, and concentration properties of (super)martingales. This paper is part of a programme initiated in [16] to study the near-critical behavior in inhomogeneous random graphs of so-called rank-1.

Key words: critical random graphs, phase transitions, inhomogeneous networks, Brownian excursions, size-biased ordering, martingale techniques.

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[†]Department of Statistics and Operations Research, University of North Carolina, 304 Hanes Hall, Chapel Hill, NC27510, United States. Email: bhamidi@email.unc.edu

[‡]Department of Mathematics and Computer Science, Eindhoven University of Technology, P.O. Box 513, 5600 MB Eindhoven, The Netherlands. E-mail: rhofstad@win.tue.nl, j.s.h.v.leeuwaarden@tue.nl

1 Introduction

1.1 Model

We start by describing the model considered in this paper. While there are many variants available in the literature, the most convenient for our purposes is the model often referred to as *Poissonian graph process* or *Norros-Reittu model* [23]. See Section 1.3 below for consequences for related models. To define the model, we consider the vertex set $[n] := \{1, 2, \dots, n\}$, and attach an edge with probability p_{ij} between vertices i and j , where

$$p_{ij} = 1 - \exp\left(-\frac{w_i w_j}{\ell_n}\right), \quad (1.1)$$

and

$$\ell_n = \sum_{i \in [n]} w_i. \quad (1.2)$$

Different edges are independent.

Below, we shall formulate general conditions on the weight sequence $\mathbf{w} = (w_i)_{i \in [n]}$, and for now formulate two main examples. The first key example arises when we take \mathbf{w} to be an i.i.d. sequence of random variables with distribution function F satisfying

$$\mathbb{E}[W^3] < \infty. \quad (1.3)$$

The second key example (which is also studied in [16]) arises when we let the weight sequence $\mathbf{w} = (w_i)_{i \in [n]}$ be defined by

$$w_i = [1 - F]^{-1}(i/n), \quad (1.4)$$

where F is a distribution function of a random variable satisfying (1.3), with $[1 - F]^{-1}$ the generalized inverse function of $1 - F$ defined, for $u \in (0, 1)$, by

$$[1 - F]^{-1}(u) = \inf\{s : [1 - F](s) \leq u\}. \quad (1.5)$$

By convention, we set $[1 - F]^{-1}(1) = 0$.

Write

$$\nu = \frac{\mathbb{E}[W^2]}{\mathbb{E}[W]}. \quad (1.6)$$

Then, by [6], the random graphs we consider are subcritical when $\nu < 1$ and supercritical when $\nu > 1$. Indeed, when $\nu > 1$, there is one giant component of size $\Theta(n)$ while all other components are of smaller size $o_{\mathbb{P}}(n)$, while when $\nu \leq 1$, the largest connected component has size $o_{\mathbb{P}}(n)$. Thus, the critical value of the model is $\nu = 1$. Here, and throughout this paper, we use the following standard notation. We write $f(n) = O(g(n))$ for functions $f, g \geq 0$ and $n \rightarrow \infty$ if there exists a constant $C > 0$ such that $f(n) \leq Cg(n)$ in the limit, and $f(n) = o(g(n))$ if $\lim_{n \rightarrow \infty} f(n)/g(n) = 0$. Furthermore, we write $f = \Theta(g)$ if $f = O(g)$ and $g = O(f)$. We write $O_{\mathbb{P}}(b_n)$ for a sequence of random variables X_n for which $|X_n|/b_n$ is tight as $n \rightarrow \infty$, and $o_{\mathbb{P}}(b_n)$ for a sequence of random variables X_n for which $|X_n|/b_n \xrightarrow{\mathbb{P}} 0$ as $n \rightarrow \infty$. Finally, we write that a sequence of events $(E_n)_{n \geq 1}$ occurs *with high probability (whp)* when $\mathbb{P}(E_n) \rightarrow 1$.

We shall write $\mathcal{G}_n^0(\mathbf{w})$ to be the graph constructed via the above procedure, while, for any fixed $t \in \mathbb{R}$, we shall write $\mathcal{G}_n^t(\mathbf{w})$ when we use the weight sequence $(1 + tn^{-1/3})\mathbf{w}$, for which the probability that i and j are neighbors equals $1 - \exp\left(-\frac{(1 + tn^{-1/3})w_i w_j}{\ell_n}\right)$. In this setting we take n so large that $1 + tn^{-1/3} > 0$.

We now formulate the general conditions on the weight sequence \mathbf{w} . In Section 3, we shall verify that these conditions are satisfied for i.i.d. weights with finite third moment, as well as for the choice in (1.4). We assume the following three conditions on the weight sequence \mathbf{w} :

(a) Maximal weight bound. We assume that the maximal weight is $o(n^{1/3})$, i.e.,

$$\max_{i \in [n]} w_i = o(n^{1/3}). \quad (1.7)$$

(b) Weak convergence of weight distribution. We assume that the weight of a uniformly chosen vertex converges in distribution to some distribution function F , i.e., let $V_n \in [n]$ be a uniformly chosen vertex. Then we assume that

$$w_{V_n} \xrightarrow{d} W, \quad (1.8)$$

for some limiting random variable W with distribution function F . Condition (1.8) is equivalent to the statement that, for every x that is a continuity point of $x \mapsto F(x)$, we have

$$\frac{1}{n} \#\{i : w_i \leq x\} \rightarrow F(x). \quad (1.9)$$

(c) Convergence of first three moments. We assume that

$$\frac{1}{n} \sum_{i \in [n]} w_i = \mathbb{E}[W] + o(n^{-1/3}), \quad (1.10)$$

$$\frac{1}{n} \sum_{i \in [n]} w_i^2 = \mathbb{E}[W^2] + o(n^{-1/3}), \quad (1.11)$$

$$\frac{1}{n} \sum_{i \in [n]} w_i^3 = \mathbb{E}[W^3] + o(1). \quad (1.12)$$

Note that condition (a) follows from conditions (b) and (c), as we prove around (2.41) below. We nevertheless choose to introduce the weaker condition (a), for its clear combinatorial interpretation and the fact that this maximal weight bounds occurs naturally at several places in the proofs. When \mathbf{w} is *random*, for example in the case where $(w_i)_{i \in [n]}$ are i.i.d. random variables with finite third moment, then we need the estimates in conditions (a), (b) and (c) to hold *in probability*.

We shall simply refer to the above three conditions as conditions (a), (b) and (c). Note that (1.10) and (1.11) in condition (c) also imply that

$$v_n = \frac{\sum_{i \in [n]} w_i^2}{\sum_{i \in [n]} w_i} = \frac{\mathbb{E}[W^2]}{\mathbb{E}[W]} + o(n^{-1/3}) = v + o(n^{-1/3}). \quad (1.13)$$

Before we write our main result we shall need one more construct. For fixed $t \in \mathbb{R}$, consider the inhomogeneous Brownian motion $(W^t(s))_{s \geq 0}$ defined as

$$W^t(s) = B(s) + st - \frac{s^2}{2}, \quad (1.14)$$

where B is standard Brownian motion, so that W^t has drift $t - s$ at time s . We want to consider this process restricted to be non-negative, and we therefore introduce the reflected process

$$\bar{W}^t(s) = W^t(s) - \min_{0 \leq s' \leq s} W^t(s'). \quad (1.15)$$

In [1] it is shown that the excursions of \bar{W}^t from 0 can be ranked in increasing order as, say, $\gamma_1(t) > \gamma_2(t) > \dots$

Now let $\mathcal{C}_n^1(t) \geq \mathcal{C}_n^2(t) \geq \mathcal{C}_n^3(t) \dots$ denote the sizes of the components in $\mathcal{G}_n^t(\mathbf{w})$ arranged in increasing order. Define l^2 to be the set of infinite sequences $x = (x_i)_{i=1}^\infty$ with $x_1 \geq x_2 \geq \dots \geq 0$ and $\sum_{i=1}^\infty x_i^2 < \infty$, and define the l^2 metric by

$$d(x, y) = \left(\sum_{i=1}^\infty (x_i - y_i)^2 \right)^{1/2}. \quad (1.16)$$

Let

$$\mu = \mathbb{E}[W], \quad \sigma_3 = \mathbb{E}[W^3]. \quad (1.17)$$

Then, our main result is as follows:

Theorem 1.1 (The critical behavior). *Assume that the weight sequence \mathbf{w} satisfies conditions (a), (b) and (c), and let $\nu = 1$. Then, as $n \rightarrow \infty$,*

$$(n^{-2/3} \mathcal{C}_n^i(t))_{i \geq 1} \xrightarrow{d} (\mu \sigma_3^{-1/3} \gamma_i(t \mu \sigma_3^{-2/3}))_{i \geq 1} =: (\gamma_i^*(t))_{i \geq 1}, \quad (1.18)$$

in distribution and with respect to the l^2 topology.

Theorem 1.1 extends the work of Aldous [1], who identified the scaling limit of the largest connected components in the Erdős-Rényi random graph. Indeed, he proved for the critical Erdős-Rényi random graph with $p = (1 + tn^{-1/3})/n$ that the ordered connected components are described by $(\gamma_i(t))_{i \geq 1}$, i.e., the ordered excursions of the reflected process in (1.15). Hence, Aldous' result corresponds to Theorem 1.1 with $\mu = \sigma_3 = 1$. The sequence $(\gamma_i^*(t))_{i \geq 1}$ is in fact the sequence of ordered excursions of the reflected version of the process

$$W_*^t(s) = \sqrt{\frac{\sigma_3}{\mu}} B(s) + st - \frac{s^2 \sigma_3}{2\mu^2}, \quad (1.19)$$

which reduces to the process in (1.14) again when $\mu = \sigma_3 = 1$.

We next investigate our two key examples, and show that conditions (a), (b) and (c) indeed hold in this case.

Corollary 1.2 (Theorem 1.1 holds for key examples). *Conditions (a), (b) and (c) are satisfied when \mathbf{w} is as in (1.4), where F is a distribution function of a random variable W with $\mathbb{E}[W^3] < \infty$, or when \mathbf{w} consists of i.i.d. copies of a random variable W with $\mathbb{E}[W^3] < \infty$.*

Theorem 1.1 was already conjectured in [16], for the case where \mathbf{w} is as in (1.4) and F is a distribution function of a random variable W with $\mathbb{E}[W^{3+\varepsilon}] < \infty$ for some $\varepsilon > 0$. The current result implies that $\mathbb{E}[W^3] < \infty$ is a sufficient condition for this result to hold, and we believe this condition also to be necessary (as the constant $\mathbb{E}[W^3]$ also appears in our results, see (1.18) and (1.19)).

1.2 Overview of the proof

In this section, we give an overview of the proof of Theorem 1.1. After having set the stage for the proof, we shall provide a heuristic that indicates how our main result comes about. We start by describing the cluster exploration:

Cluster exploration. The proof involves two key ingredients:

- The exploration of components via breath-first search; and
- The labeling of vertices in a size-biased order of their weights \mathbf{w} .

More precisely, we shall explore components and simultaneously construct the graph $\mathcal{G}_n^t(\mathbf{w})$ in the following manner: First, for all ordered pairs of vertices (i, j) , let $V(i, j)$ be exponential random variables with rate $(1 + tn^{-1/3})w_j/\ell_n$. Choose vertex $v(1)$ with probability proportional to \mathbf{w} , so that

$$\mathbb{P}(v(1) = i) = w_i/\ell_n. \quad (1.20)$$

The children of $v(1)$ are all the vertices j such that

$$V(v(1), j) \leq w_{v(1)}. \quad (1.21)$$

Suppose $v(1)$ has $c(1)$ children. Label these as $v(2), v(3), \dots, v(c(1) + 1)$ in increasing order of their $V(v(1), \cdot)$ values. Now move on to $v(2)$ and explore all of its children (say $c(2)$ of them) and label them as before. Note that when we explore the children of $v(2)$, its potential children cannot include the vertices that we have already identified. More precisely, the children of $v(2)$ consist of the set

$$\{v \notin \{v(1), \dots, v(c(1) + 1)\} : V(v(2), v) \leq w_{v(2)}\}$$

and so on. Once we finish exploring one component, we move on to the next component by choosing the starting vertex in a size-biased manner amongst the remaining vertices and start exploring its component. It is obvious that in this way we find all the components of our graph $\mathcal{G}_n^t(\mathbf{w})$.

Write the breadth-first walk associated to this exploration process as

$$Z_n(0) = 0, \quad Z_n(i) = Z_n(i - 1) + c(i) - 1, \quad (1.22)$$

for $i = 1, \dots, n$. Suppose $\mathcal{C}^*(i)$ is the size of the i^{th} component explored in this manner (here we write $\mathcal{C}^*(i)$ to distinguish this from $\mathcal{C}_n^i(t)$, the i^{th} largest component). Then these can be easily recovered from the above walk by the following prescription: For $j \geq 0$, write $\eta(j)$ as the stopping time

$$\eta(j) = \min\{i : Z_n(i) = -j\}. \quad (1.23)$$

Then

$$\mathcal{C}^*(j) = \eta(j) - \eta(j - 1). \quad (1.24)$$

Further,

$$Z_n(\eta(j)) = -j, \quad Z_n(i) \geq -j \text{ for all } \eta(j) < i < \eta(j + 1). \quad (1.25)$$

Recall that we started with vertices labeled $1, 2, \dots, n$ with corresponding weights $\mathbf{w} = (w_i)_{i \in [n]}$. The size-biased order $v^*(1), v^*(2), \dots, v^*(n)$ is a random reordering of the above vertex set where $v(1) = i$ with probability equal to w_i/ℓ_n . Then, given $v^*(1)$, we have that $v^*(2) = j \in [n] \setminus \{v^*(1)\}$ with probability proportional to w_j and so on. By construction and the properties of the exponential random variables, we have the following representation, which lies at the heart of our analysis:

Lemma 1.3 (Size-biased reordering of vertices). *The order $v(1), v(2), \dots, v(n)$ in the above construction of the breadth-first exploration process is the size-biased ordering $v^*(1), v^*(2), \dots, v^*(n)$ of the vertex set $[n]$ with weights proportional to w .*

Proof. The first vertex $v(1)$ is chosen from $[n]$ via the size-biased distribution. Suppose it has no neighbors. Then, by construction, the next vertex is chosen via the size-biased distribution amongst all remaining vertices. If vertex 1 does have neighbors, then we shall use the following construction. For $j \geq 2$, choose $V(v(1), j)$ exponentially distributed with rate $(1 + tn^{-1/3})w_j/\ell_n$. Rearrange the vertices in increasing order of their $V(v(1), j)$ values (so that $v'(2)$ is the vertex with the smallest $V(v(1), j)$ value, $v'(3)$ is the vertex with the second smallest value and so on). Note that by the properties of the exponential distribution

$$\mathbb{P}(v(2) = i \mid v(1)) = \frac{w_i}{\sum_{j \neq v(1)} w_j} \quad \text{for } j \in [n] \setminus \{v(1)\}. \quad (1.26)$$

Similarly, given the value of $v(2)$,

$$\mathbb{P}(v(3) = i \mid v(1), v(2)) = \frac{w_i}{\sum_{j \neq v(1), v(2)} w_j}, \quad (1.27)$$

and so on. Thus the above gives us a size-biased ordering of the vertex set $[n] \setminus \{v(1)\}$. Suppose $c(1)$ of the exponential random variables are less than $w_{v(1)}$. Then set $v(j) = v'(j)$ for $2 \leq j \leq c(1) + 1$ and discard all the other labels. This gives us the first $c(1) + 1$ values of our size-biased ordering.

Once we are done with $v(1)$, let the potentially unexplored neighbors of $v(2)$ be

$$\mathcal{U}_2 = [n] \setminus \{v(1), \dots, v(c(1) + 1)\}, \quad (1.28)$$

and, again, for j in \mathcal{U}_2 , we let $V(v(2), j)$ be exponential with rate $(1 + tn^{-1/3})w_j/\ell_n$ and proceed as above.

Proceeding this way, it is clear that at the end, the random ordering $v(1), v(2), \dots, v(n)$ that we obtain is a size-biased random ordering of the vertex set $[n]$. This proves the lemma. \square

Heuristic derivation of Theorem 1.1. We next provide a heuristic that explains the limiting process in (1.19). Note that by our assumptions on the weight sequence we have for the graph $\mathcal{G}_n^t(w)$

$$p_{ij} = \left(1 + o(n^{-1/3})\right) p_{ij}^*, \quad (1.29)$$

where

$$p_{ij}^* = \left(1 + tn^{-1/3}\right) \frac{w_i w_j}{\ell_n}. \quad (1.30)$$

In the remainder of the proof, wherever we need p_{ij} , we shall use p_{ij}^* instead, which shall simplify the calculations and exposition.

Recall the cluster exploration described above, and, in particular, Lemma 1.3. We explore the cluster one vertex at a time, in a breadth-first manner. We choose $v(1)$ according to w , i.e., $\mathbb{P}(v(1) = j) = w_j/\ell_n$. We say that a vertex is *explored* when its neighbors have been investigated, and *unexplored* when it has been found to be part of the cluster found so far, but its neighbors have not been

investigated yet. Finally, we say that a vertex is *neutral* when it has not been considered at all. Thus, in our cluster exploration, as long as there are unexplored vertices, we explore the vertices $(v(i))_{i \in [n]}$ in the order of appearance. When there are no unexplored vertices left, then we draw (size-biased) from the neutral vertices. Then, Lemma 1.3 states that $(v(i))_{i \in [n]}$ is a size-biased reordering of $[n]$.

Let $c(i)$ denote the number of neutral neighbors of $v(i)$, and recall (1.22). The clusters of our random graph are found in between successive times in which $(Z_n(l))_{l \in [n]}$ reaches a new minimum. Now, Theorem 1.1 follows from the fact that $\bar{Z}_n(s) = n^{-1/3} Z_n(\lfloor n^{2/3}s \rfloor)$ weakly converges to $(W_*^t(s))_{s \geq 0}$ defined as in (1.19). General techniques from [1] show that this also implies that the ordered excursions between successive minima of $(\bar{Z}_n(s))_{s \geq 0}$ converge to the ones of $(W_*^t(s))_{s \geq 0}$. These ordered excursions were denoted by $\gamma_1^*(t) > \gamma_2^*(t) > \dots$. Using Brownian scaling, it can be seen that

$$W_*^t(s) \stackrel{d}{=} \sigma_3^{1/3} W^t \mu \sigma_3^{-2/3} (\sigma_3^{1/3} \mu^{-1} s) \quad (1.31)$$

with W_t defined in (1.14). Hence, from the relation (1.31) it immediately follows that

$$(\gamma_i^*(t))_{i \geq 1} \stackrel{d}{=} (\mu \sigma_3^{-1/3} \gamma_i(t \mu \sigma_3^{-2/3}))_{i \geq 1}, \quad (1.32)$$

which then proves Theorem 1.1.

To see how to derive (1.31), fix $a > 0$ and note that $(B(a^2s))_{s \geq 0}$ has the same distribution as $(aB(s))_{s \geq 0}$. Thus, for $(W_{\sigma, \kappa}^t(s))_{s \geq 0}$ with

$$W_{\sigma, \kappa}^t(s) = \sigma B(s) + st - \kappa s^2/2, \quad (1.33)$$

we obtain the scaling relation

$$W_{\sigma, \kappa}^t(s) \stackrel{d}{=} \frac{\sigma}{a} W_{1, a^{-3}\kappa/\sigma}^t(a^2s). \quad (1.34)$$

Using $\kappa = \sigma^2/\mu$ and $a = (\kappa/\sigma)^{1/3} = (\sigma/\mu)^{1/3}$, we note that

$$W_{\sigma, \sigma^2/\mu}^t(s) \stackrel{d}{=} \sigma^{2/3} \mu^{1/3} W^{t\sigma^{-4/3} \mu^{1/3}}(\sigma^{2/3} \mu^{-2/3} s), \quad (1.35)$$

which, with $\sigma = (\sigma_3/\mu)^{1/2}$, yields (1.31).

We complete the sketch of proof by giving a heuristic argument that indeed $\bar{Z}_n(s) = n^{-1/3} Z_n(\lfloor n^{2/3}s \rfloor)$ weakly converges to $(W_*^t(s))_{s \geq 0}$. For this, we investigate $c(i)$, the number of neutral neighbors of $v(i)$. Throughout this paper, we shall denote

$$\tilde{w}_j = w_j(1 + tn^{-1/3}), \quad (1.36)$$

so that the $\mathcal{G}_n^t(\mathbf{w})$ has weights $\tilde{\mathbf{w}} = (\tilde{w}_j)_{j \in [n]}$.

We note that since p_{ij} in (1.1) is quite small, the number of neighbors of a vertex j is close to $\text{Poi}(\tilde{w}_j)$, where $\text{Poi}(\lambda)$ denotes a Poisson random variable with mean λ . Thus, the number of neutral neighbors is close to the total number of neighbors minus the active neighbors, i.e.,

$$c(i) \approx \text{Poi}(\tilde{w}_{v(i)}) - \text{Poi}\left(\sum_{j=1}^{i-1} \frac{\tilde{w}_{v(i)} \tilde{w}_{v(j)}}{\ell_n}\right), \quad (1.37)$$

since $\sum_{j=1}^{i-1} \tilde{w}_{v(i)} \tilde{w}_{v(j)} / \ell_n$ is, conditionally on $(v(j))_{j=1}^i$, the expected number of edges between $v(i)$ and $(v(j))_{j=1}^{i-1}$. We conclude that the increase of the process $Z_n(l)$ equals

$$c(i) - 1 \approx \text{Poi}(\tilde{w}_{v(i)}) - 1 - \text{Poi}\left(\sum_{j=1}^{i-1} \frac{\tilde{w}_{v(i)} \tilde{w}_{v(j)}}{\ell_n}\right), \quad (1.38)$$

so that

$$Z_n(l) \approx \sum_{i=1}^l (\text{Poi}(\tilde{w}_{v(i)}) - 1) - \text{Poi}\left(\sum_{j=1}^{i-1} \frac{\tilde{w}_{v(i)} \tilde{w}_{v(j)}}{\ell_n}\right). \quad (1.39)$$

The change in $Z_n(l)$ is not stationary, and decreases on the average as l increases, due to two reasons. First of all, the number of neutral vertices decreases (as is apparent from the sum which is subtracted in (1.39)), and the law of $\tilde{w}_{v(l)}$ becomes stochastically smaller as l increases. The latter can be understood by noting that $\mathbb{E}[\tilde{w}_{v(1)}] = (1 + tn^{-1/3})\nu_n = 1 + tn^{-1/3} + o(n^{-1/3})$, while $\frac{1}{n} \sum_{j \in [n]} \tilde{w}_{v(j)} = (1 + tn^{-1/3})\ell_n/n$, and, by Cauchy-Schwarz,

$$\ell_n/n \approx \mathbb{E}[W] \leq \mathbb{E}[W^2]^{1/2} = \mathbb{E}[W]^{1/2} \nu^{1/2} = \mathbb{E}[W]^{1/2}, \quad (1.40)$$

so that $\ell_n/n \leq 1 + o(1)$, and the inequality becomes strict when $\text{Var}(W) > 0$. We now study these two effects in more detail.

The random variable $\text{Poi}(\tilde{w}_{v(i)}) - 1$ has asymptotic mean

$$\mathbb{E}[\text{Poi}(\tilde{w}_{v(i)}) - 1] \approx \sum_{j \in [n]} \tilde{w}_j \mathbb{P}(v(i) = j) - 1 \approx \sum_{j \in [n]} \tilde{w}_j \frac{w_j}{\ell_n} - 1 = \nu_n(1 + tn^{-1/3}) - 1 \approx 0. \quad (1.41)$$

However, since we sum $\Theta(n^{2/3})$ contributions, and we multiply by $n^{-1/3}$, we need to be rather precise and compute error terms up to order $n^{-1/3}$ in the above computation. We shall do this now, by conditioning on $(v(j))_{j=1}^{i-1}$. Indeed, writing $\mathbb{1}_A$ for the indicator of the event A ,

$$\begin{aligned} \mathbb{E}[\tilde{w}_{v(i)} - 1] &\approx \nu_n tn^{-1/3} + \mathbb{E}[\mathbb{E}[w_{v(i)} - 1 \mid (v(j))_{j=1}^{i-1}]] \\ &\approx tn^{-1/3} + \mathbb{E}\left[\sum_{l=1}^n w_l \mathbb{1}_{\{l \notin \{v(j)\}_{j=1}^{i-1}\}} \frac{w_l}{\ell_n - \sum_{j=1}^{i-1} w_{v(j)}}\right] - 1 \\ &\approx tn^{-1/3} + \sum_{j \in [n]} \frac{w_j^2}{\ell_n} + \mathbb{E}\left[\frac{1}{\ell_n^2} \sum_{j=1}^{i-1} w_{v(j)} \sum_{l=1}^n w_l^2\right] - \mathbb{E}\left[\frac{1}{\ell_n} \sum_{j=1}^{i-1} w_{v(j)}^2\right] - 1 \\ &\approx tn^{-1/3} + i\left(\frac{\nu_n^2}{\ell_n} - \frac{1}{\ell_n^2} \sum_{j \in [n]} w_j^3\right) \approx tn^{-1/3} + \frac{i}{\ell_n} \left(1 - \frac{1}{\ell_n} \sum_{j \in [n]} w_j^3\right). \end{aligned} \quad (1.42)$$

When $i = \Theta(n^{2/3})$, these terms are indeed both of order $n^{-1/3}$, and shall thus contribute to the scaling limit of $(Z_n(l))_{l \geq 0}$.

The variance of $\text{Poi}(\tilde{w}_{v(i)})$ is approximately equal to

$$\begin{aligned} \text{Var}(\text{Poi}(\tilde{w}_{v(i)})) &= \mathbb{E}[\text{Var}(\text{Poi}(\tilde{w}_{v(i)})) \mid v(i)] + \text{Var}(\mathbb{E}[\text{Poi}(\tilde{w}_{v(i)}) \mid v(i)]) \\ &= \mathbb{E}[\tilde{w}_{v(i)}] + \text{Var}(\tilde{w}_{v(i)}) \approx \mathbb{E}[\tilde{w}_{v(i)}^2] \approx \mathbb{E}[w_{v(i)}^2], \end{aligned} \quad (1.43)$$

since $\mathbb{E}[w_{v(i)}] = 1 + \Theta(n^{-1/3})$. Summing the above over $i = 1, \dots, sn^{2/3}$ and multiplying by $n^{-1/3}$ intuitively explains that

$$n^{-1/3} \sum_{i=1}^{sn^{2/3}} (\text{Poi}(\tilde{w}_{v(i)}) - 1) \xrightarrow{d} \sigma B(s) + st + \frac{s^2}{2\mathbb{E}[W]}(1 - \sigma^2), \quad (1.44)$$

where we write $\sigma^2 = \mathbb{E}[W^3]/\mathbb{E}[W]$ and we let $(B(s))_{s \geq 0}$ denote a standard Brownian motion. We also adopt the convention that when a non-integer, such as $sn^{2/3}$ appears in summation bounds, it should be rounded down. Note that when $\text{Var}(W) > 0$, then $\mathbb{E}[W] = \mathbb{E}[W^2] < 1$, so that $\mathbb{E}[W^3]/\mathbb{E}[W] > 1$ and the constant in front of s^2 is *negative*. We shall make the limit in (1.44) precise by using a *martingale functional central limit theorem*.

The second term in (1.39) turns out to be well-concentrated around its mean, so that, in this heuristic, we shall replace it by its mean. The concentration shall be proved using concentration techniques on appropriate supermartingales. This leads us to compute

$$\begin{aligned} \mathbb{E} \left[\sum_{i=1}^l \text{Poi} \left(\sum_{j=1}^{i-1} \frac{\tilde{w}_{v(i)} \tilde{w}_{v(j)}}{\ell_n} \right) \right] &\approx \mathbb{E} \left[\sum_{i=1}^l \sum_{j=1}^{i-1} \frac{\tilde{w}_{v(i)} \tilde{w}_{v(j)}}{\ell_n} \right] \approx \mathbb{E} \left[\sum_{i=1}^l \sum_{j=1}^{i-1} \frac{w_{v(i)} w_{v(j)}}{\ell_n} \right] \\ &\approx \frac{1}{2} \mathbb{E} \left[\frac{1}{\ell_n} \left(\sum_{j=1}^l w_{v(j)} \right)^2 \right] \approx \frac{1}{2\ell_n} \mathbb{E} \left[\sum_{j=1}^l w_{v(j)} \right]^2, \end{aligned} \quad (1.45)$$

the last asymptotic equality again following from the fact that the random variable involved is concentrated.

We conclude that

$$n^{-1/3} \mathbb{E} \left[\sum_{i=1}^{sn^{2/3}} \text{Poi} \left(\sum_{j=1}^{i-1} \frac{\tilde{w}_{v(i)} \tilde{w}_{v(j)}}{\ell_n} \right) \right] \approx \frac{s^2}{2\mathbb{E}[W]}. \quad (1.46)$$

Subtracting (1.46) from (1.44), these computations suggest, informally, that

$$\begin{aligned} \bar{Z}_n(s) = n^{-1/3} Z_n(\lfloor n^{2/3}s \rfloor) &\xrightarrow{d} \sigma B(s) + st - \frac{s^2 \mathbb{E}[W^3]}{2\mathbb{E}[W]^2} \\ &= \sqrt{\frac{\mathbb{E}[W^3]}{\mathbb{E}[W]}} B(s) + st - \frac{s^2 \mathbb{E}[W^3]}{2\mathbb{E}[W]^2}, \end{aligned} \quad (1.47)$$

as required. Note the cancelation of the terms $s^2/(2\mathbb{E}[W])$ in (1.44) and (1.46), where they appear with an opposite sign. Our proof will make this analysis precise.

1.3 Discussion

Our results are generalizations of the critical behavior of Erdős-Rényi random graphs, which have received tremendous attention over the past decades. We refer to [1], [5], [19] and the references therein. Properties of the limiting distribution of the largest component $\gamma_1(t)$ can be found in [24], which, together with the recent local limit theorems in [17], give excellent control over the joint tail behavior of several of the largest connected components.

Comparison to results of Aldous. We have already discussed the relation between Theorem 1.1 and the results of Aldous on the largest connected components in the Erdős-Rényi random graph. However, Theorem 1.1 is related to another result of Aldous [1, Proposition 4], which is less well known, and which investigates a kind of Norros-Reittu model (see [23]) for which the ordered *weights* of the clusters are determined. Here, the weight of a set of vertices $A \subseteq [n]$ is defined by $\bar{w}_A = \sum_{a \in A} w_a$. Indeed, Aldous defines an inhomogeneous random graph where the edge probability is equal to

$$p_{ij} = 1 - e^{-q x_i x_j}, \quad (1.48)$$

and assumes that the pair $(q, (x_i)_{i \in [n]})$ satisfies the following scaling relation:

$$\frac{\sum_{i \in [n]} x_i^3}{(\sum_{i \in [n]} x_i^2)^3} \rightarrow 1, \quad q - \left(\sum_{i \in [n]} x_i^2 \right)^{-1} \rightarrow t, \quad \max_{j \in [n]} x_j = o\left(\sum_{i \in [n]} x_i^2 \right). \quad (1.49)$$

When we pick

$$x_j = w_j \frac{(\sum_{i \in [n]} w_i^3)^{1/3}}{\sum_{i \in [n]} w_i^2}, \quad q = \frac{(\sum_{i \in [n]} w_i^2)^2}{(\sum_{i \in [n]} w_i^3)^{2/3} \ell_n} (1 + t n^{-1/3}), \quad (1.50)$$

then these assumptions are very similar to conditions (a)-(c). However, the asymptotics of q in (1.49) is replaced with

$$\begin{aligned} q - \left(\sum_{i \in [n]} x_i^2 \right)^{-1} &= \frac{\frac{1}{n} \sum_{i \in [n]} w_i^2}{\left(\frac{1}{n} \sum_{i \in [n]} w_i^3 \right)^{2/3}} (n^{1/3} \nu_n (1 + t n^{-1/3}) - n^{1/3}) \\ &\rightarrow \frac{\mathbb{E}[W^2]}{\mathbb{E}[W^3]^{2/3}} t = \frac{\mathbb{E}[W]}{\mathbb{E}[W^3]^{2/3}} t, \end{aligned} \quad (1.51)$$

where the last equality follows from the fact that $\nu = \mathbb{E}[W^2]/\mathbb{E}[W] = 1$. This scaling in t simply means that the parameter t in the process $W_*^t(s)$ in (1.19) is rescaled, which is explained in more detail in the scaling relations in (1.32). Write $\mathcal{C}_n^i(t)$ for the component with the i^{th} largest **weight**, and let $\bar{w}_{\mathcal{C}_n^i(t)} = \sum_{j \in \mathcal{C}_n^i(t)} w_j$ denote the cluster weight. Then, Aldous [1, Proposition 4] proves that

$$\left(\frac{(\sum_{i \in [n]} w_i^3)^{1/3}}{\sum_{i \in [n]} w_i^2} \bar{w}_{\mathcal{C}_n^i(t)} \right)_{i \geq 1} \xrightarrow{d} (\gamma_i(t \mathbb{E}[W]/\mathbb{E}[W^3]^{2/3}))_{i \geq 1}, \quad (1.52)$$

where we recall that $(\gamma_i(t))_{i \geq 1}$ is the scaling limit of the ordered component sizes in the Erdős-Rényi random graph with parameter $p = (1 + t n^{-1/3})/n$. Now,

$$\frac{(\sum_{i \in [n]} w_i^3)^{1/3}}{\sum_{i \in [n]} w_i^2} \approx n^{-2/3} \mathbb{E}[W^3]^{1/3} / \mathbb{E}[W^2] = n^{-2/3} \mathbb{E}[W^3]^{1/3} / \mathbb{E}[W], \quad (1.53)$$

and one would expect that $\bar{w}_{\mathcal{C}_n^i(t)} \approx \mathcal{C}_n^i(t)$, which is consistent with (1.18) and (1.32). The technique used by Aldous [1] to deal with the *ordered cluster weights* does not apply immediately to our setting of ordered component sizes, as Aldous [1] relies on a continuous-time description of the cluster weight exploration. As a result, we use slightly adapted (super)martingale techniques.

Related models. The model studied here is asymptotically equivalent to many related models appearing in the literature, for example to the *random graph with prescribed expected degrees* that has been studied intensively by Chung and Lu (see [8, 9, 10, 11, 12]). This model corresponds to the rank-1 case of the general inhomogeneous random graphs studied in [6] and satisfies

$$p_{ij} = \min \left\{ \frac{w_i w_j}{\ell_n}, 1 \right\}. \quad (1.54)$$

Further, for the *generalized random graph* introduced by Britton, Deijfen and Martin-Löf in [7], the edge occupation probabilities are given by

$$p_{ij} = \frac{w_i w_j}{\ell_n + w_i w_j}. \quad (1.55)$$

See [16, 18] for more details on the asymptotic equivalence of such inhomogeneous random graphs. Further, Nachmias and Peres [22] recently proved similar scaling limits for critical percolation on random regular graphs.

Alternative approach by Turova. Turova [26] recently obtained results for a setting that is similar to ours. Turova takes the edge probabilities to be $p_{ij} = \min\{x_i x_j / n, 1\}$, and assumes that $(x_i)_{i \in [n]}$ are i.i.d. random variables with $\mathbb{E}[X^3] < \infty$. This setting follows from ours by taking

$$w_i = x_i \left(\frac{1}{n} \sum_{j \in [n]} x_j \right). \quad (1.56)$$

Naturally, the critical point changes in Turova's setting, and becomes $\mathbb{E}[X^2] = 1$.

First versions of the paper [26] and this paper were uploaded almost simultaneously on the ArXiv. Comparing the two papers gives interesting insights in how to deal with the inherent size-biased orderings in two rather different ways. Turova applies discrete martingale techniques in the spirit of Martin-Löf's [21] work on diffusion approximations for critical epidemics, while our approach is more along the lines of the original paper of Aldous [1], relying on concentration techniques and supermartingales (see Lemma 2.2). Further, our result is slightly more general than the one in [26]. In fact, our discussions with Turova inspired us to extend our setting to one that includes i.i.d. weights (which is Turova's original setting).

The necessity of conditions (a)-(c). The conditions (a)-(c) provide conditions under which we prove convergence. One may wonder whether these conditions are merely sufficient, or also necessary. Condition (b) gives stability of the weight structure, which implies that the local neighborhoods in our random graphs locally converge to appropriate branching processes. The latter is a strengthening of the assumption that our random graphs are *sparse*, and is a natural condition to start with. We believe that, given that condition (b) holds, conditions (a) and (c) are necessary. Indeed, Aldous and Limic give several examples where the scaling of the largest critical cluster is $n^{2/3}$ with a *different* scaling limit when $w_1 n^{1/3} \rightarrow c_1 > 0$ (see [2, Proof of lemma 8, p. 10]). Therefore, for Theorem 1.1 to hold (with the prescribed scaling limit in terms of ordered Brownian excursions), condition (a) seems to be necessary. Since conditions (b) and (c) imply condition (a), it follows that if we assume condition (b), then we need the other two conditions for our main result to hold. This answers [1, Open problem 2, p. 851].

Inhomogeneous random graphs with infinite third moments. In the present setting, when it is assumed that $\mathbb{E}[W^3] < \infty$, the scaling limit turns out to be a scaled version of the scaling limit for the Erdős-Rényi random graph as identified in [1]. In [3], we have recently studied the case where $\mathbb{E}[W^3] = \infty$, for which the critical behavior turns out to be fundamentally different. More specifically, we choose in [3] the weight sequence $\mathbf{w} = (w_i)_{i \in [n]}$ as in (1.4) with F such that, as $x \rightarrow \infty$, $1 - F(x) = c_F x^{-(\tau-1)}(1 + o(1))$, for some $\tau \in (3, 4)$ and $0 < c_F < \infty$. Under this assumption, the clusters have asymptotic size $n^{(\tau-2)/(\tau-1)}$ (see [16]). The scaling limit itself turns out to be described in terms of a so-called ‘thinned’ Lévy process, that consists of infinitely many Poisson processes with varying rates of which only the first event is counted, and which already appeared in [2] in the context of random graphs having $n^{2/3}$ critical behavior. Moreover, we prove in [3] that the vertex i is in the largest connected component with non-vanishing probability as $n \rightarrow \infty$, which implies that the highest weight vertices characterize the largest components (‘power to the wealthy’). This is in sharp contrast to the present setting, where the probability that the vertex with the largest weight is in the largest component is negligible, and instead the largest connected component is an extreme value event arising from many trials with *roughly* equal probability (‘power to the masses’).

2 Weak convergence of cluster exploration

In this section, we shall study the scaling limit of the cluster exploration studied in Section 1.2 above. The main result in this paper, in its most general form, is the following theorem:

Theorem 2.1 (Weak convergence of cluster exploration). *Assume that the weight sequence \mathbf{w} satisfies conditions (a), (b) and (c), and that $\nu = 1$. Consider the breadth-first walk $Z_n(\cdot)$ of (1.25) exploring the components of the random graph $\mathcal{G}_n^t(\mathbf{w})$. Define*

$$\bar{Z}_n(s) = n^{-1/3} Z_n(\lfloor n^{2/3}s \rfloor). \quad (2.1)$$

Then, as $n \rightarrow \infty$,

$$\bar{Z}_n \xrightarrow{d} W_*^t, \quad (2.2)$$

where W_*^t is the process defined in (1.19), in the sense of convergence in the J_1 Skorohod topology on the space of right-continuous left-limited functions on \mathbb{R}^+ .

To show how Theorem 2.1 immediately proves Theorem 1.1, we compare (1.15) and (1.25). Theorem 2.1 suggests that also the excursions of \bar{Z}_n beyond past minima arranged in increasing order converge to the corresponding excursions of W_*^t beyond past minima arranged in increasing order. See Aldous [1, Section 3.3] for a proof of this fact. Therefore, Theorem 2.1 implies Theorem 1.1. The remainder of this paper is devoted to the proof of Theorem 2.1.

Proof of Theorem 2.1. We shall make use of a martingale central limit theorem. By (1.29),

$$p_{ij} \approx \left(1 + \frac{t}{n^{1/3}}\right) \frac{w_i w_j}{\ell_n}, \quad (2.3)$$

and we shall use the above as an equality for the rest of the proof as this shall simplify exposition. It is quite easy to show that the error made is negligible in the limit.

Recall from (1.22) that

$$Z_n(k) = \sum_{i=1}^k (c(i) - 1). \quad (2.4)$$

Then, we decompose

$$Z_n(k) = M_n(k) + A_n(k), \quad (2.5)$$

where

$$M_n(k) = \sum_{i=1}^k (c(i) - \mathbb{E}[c(i) | \mathcal{F}_{i-1}]), \quad A_n(k) = \sum_{i=1}^k \mathbb{E}[c(i) - 1 | \mathcal{F}_{i-1}], \quad (2.6)$$

with \mathcal{F}_i the natural filtration of Z_n . Then, clearly, $\{M_n(k)\}_{k=0}^n$ is a martingale. For a process $\{S_k\}_{k=0}^n$, we further write

$$\bar{S}_n(u) = n^{-1/3} S_n(\lfloor un^{2/3} \rfloor). \quad (2.7)$$

Furthermore, let

$$B_n(k) = \sum_{i=1}^k (\mathbb{E}[c(i)^2 | \mathcal{F}_{i-1}] - \mathbb{E}[c(i) | \mathcal{F}_{i-1}]^2). \quad (2.8)$$

Then, by the martingale central limit theorem ([15, Theorem 7.1.4]), Theorem 2.1 follows when the following three conditions hold:

$$\sup_{s \leq u} \left| \bar{A}_n(s) + \frac{s^2 \sigma_3}{2\mu^2} - st \right| \xrightarrow{\mathbb{P}} 0, \quad (2.9)$$

$$n^{-2/3} B_n(n^{2/3}u) \xrightarrow{\mathbb{P}} \frac{\sigma_3 u}{\mu}, \quad (2.10)$$

$$\mathbb{E}(\sup_{s \leq u} |\bar{M}_n(s) - \bar{M}_n(s-)|^2) \longrightarrow 0. \quad (2.11)$$

Indeed, the last two equations, by [15, Theorem 7.1.4] imply that the process $\bar{M}_n(s) = n^{-1/3} M_n(n^{2/3}s)$ satisfies the asymptotics

$$\bar{M}_n \xrightarrow{d} \sqrt{\frac{\sigma_3}{\mu}} B, \quad (2.12)$$

where, as before, B is a standard Brownian motion, while (2.9) gives the drift term in (1.19) and this completes the proof.

We shall now start to verify the conditions (2.9), (2.10) and (2.11). Throughout the proof, we shall assume, without loss of generality, that $w_1 \geq w_2 \geq \dots \geq w_n$. Recall that we shall work with weight sequence $\tilde{w} = (1 + tn^{-1/3})\mathbf{w}$, for which the edge probabilities are approximately equal to $w_i w_j (1 + tn^{-1/3}) / \ell_n$ (recall (2.3)).

We note that, since $M_n(k)$ is a discrete martingale,

$$\begin{aligned} \sup_{s \leq u} |\bar{M}_n(s) - \bar{M}_n(s-)|^2 &= n^{-2/3} \sup_{k \leq un^{2/3}} (M_n(k) - M_n(k-1))^2 \leq n^{-2/3} (1 + \sup_{k \leq un^{2/3}} c(k)^2) \\ &\leq n^{-2/3} (1 + \Delta_n^2), \end{aligned} \quad (2.13)$$

where Δ_n is the maximal degree in the graph, so that

$$\mathbb{E}(\sup_{s \leq u} |\bar{M}_n(s) - \bar{M}_n(s-)|^2) \leq n^{-2/3} (1 + \mathbb{E}[\Delta_n^2]). \quad (2.14)$$

By condition (a), $\tilde{w}_i = o(n^{1/3})$, so that, by Cauchy-Schwarz,

$$\mathbb{E}[\Delta_n^2] \leq \mathbb{E}[\Delta_n^4]^{1/2} \leq n^{1/2} \left(\frac{1}{n} \sum_{i \in [n]} \mathbb{E}[d_i^4] \right)^{1/2} \leq c\sqrt{n} \left(\frac{1}{n} \sum_{i \in [n]} (w_i^4 + w_i) \right)^{1/2} = o(n^{2/3}), \quad (2.15)$$

where we use the fact that $\mathbb{E}[\text{Poi}(\lambda)^4] = O(\lambda^4 + \lambda)$ for all $\lambda > 0$. This proves that the r.h.s. of (2.14) is $o(1)$ and thus proves (2.11).

We continue with (2.9) and (2.10), for which we first analyse $c(i)$. In the course of the proof, we shall make use of the following lemma, which lies at the core of the argument:

Lemma 2.2 (Sums over sized-biased orderings). *As $n \rightarrow \infty$, for all $t > 0$,*

$$\sup_{u \leq t} \left| n^{-2/3} \sum_{i=1}^{n^{2/3}u} w_{v(i)}^2 - \frac{\sigma_3 u}{\mu} \right| \xrightarrow{\mathbb{P}} 0, \quad (2.16)$$

$$n^{-2/3} \sum_{i=1}^{n^{2/3}u} \mathbb{E}[w_{v(i)}^2 \mid \mathcal{F}_{i-1}] \xrightarrow{\mathbb{P}} \frac{\sigma_3 u}{\mu}. \quad (2.17)$$

Proof. We start by proving (2.16), for which we write

$$H_n(u) = n^{-2/3} \sum_{i=1}^{\lfloor un^{2/3} \rfloor} w_{v(i)}^2. \quad (2.18)$$

We shall use a randomization trick introduced by Aldous [1]. Let T_j be a sequence of independent exponential random variables with rate w_j/ℓ_n and define

$$\tilde{H}_n(v) = n^{-2/3} \sum_{i \in [n]} w_i^2 \mathbf{1}\{T_j \leq n^{2/3}v\}. \quad (2.19)$$

Note that by the properties of the exponential random variables, if we rank the vertices according to the order in which they arrive, then they appear in size-biased order. More precisely, for any v ,

$$\tilde{H}_n(v) = n^{-2/3} \sum_{j \in [n]} w_j^2 \mathbf{1}\{T_j \leq n^{2/3}v\} = n^{-2/3} \sum_{i=1}^{N(vn^{2/3})} w_{v(i)}^2 = H_n(N(vn^{2/3})), \quad (2.20)$$

where

$$N(t) := \#\{j : T_j \leq t\}. \quad (2.21)$$

As a result, when $N(2tn^{2/3}) \geq tn^{2/3}$ **whp**, we have that

$$\begin{aligned} \sup_{u \leq t} \left| n^{-2/3} \sum_{i=1}^{n^{2/3}u} w_{v(i)}^2 - \frac{\sigma_3 u}{\mu} \right| &\leq \sup_{u \leq 2t} \left| n^{-2/3} \sum_{i=1}^{N(un^{2/3})} w_{v(i)}^2 - \frac{\sigma_3}{\mu} n^{-2/3} N(un^{2/3}) \right| + \frac{\sigma_3 n^{-2/3}}{\mu} \\ &\leq \sup_{u \leq 2t} \left| \tilde{H}_n(u) - \frac{\sigma_3}{\mu} u \right| + \frac{\sigma_3}{\mu} \sup_{u \leq 2t} \left| n^{-2/3} N(un^{2/3}) - u \right| + \frac{\sigma_3 n^{-2/3}}{\mu}. \end{aligned} \quad (2.22)$$

The last term is due to a usual ‘rounding by one error’, and is negligible. We shall prove that the first two terms both converge to zero in probability. We start with the second, for which we use that the process

$$Y_0(s) = \frac{1}{n^{1/3}} \left(N(sn^{2/3}) - sn^{2/3} \right) \quad (2.23)$$

is a supermartingale, since

$$\begin{aligned} \mathbb{E}[N(t+s) \mid \mathcal{F}_t] &= N(t) + \mathbb{E}[N(t+s) - N(t) \mid \mathcal{F}_t] \leq N(t) + \mathbb{E}[\#\{j : T_j \in (t, t+s]\} \mid \mathcal{F}_t] \\ &\leq N(t) + \sum_{j \in [n]} (1 - e^{-w_j s / \ell_n}) \leq N(t) + \sum_{j \in [n]} \frac{w_j s}{\ell_n} = N(t) + s, \end{aligned} \quad (2.24)$$

as required. Therefore,

$$|\mathbb{E}[Y_0(t)]| = -\mathbb{E}[Y_0(t)] = \frac{1}{n^{1/3}} \left[tn^{2/3} - \sum_{i \in [n]} (1 - \exp(-tn^{2/3} w_i / \ell_n)) \right]. \quad (2.25)$$

Using the fact that $1 - e^{-x} \geq x - x^2/2$, for $x \geq 0$, we obtain that, also using the fact that $\nu_n = 1 + o(1)$,

$$|\mathbb{E}[Y_0(t)]| \leq \sum_{i \in [n]} \frac{nw_i^2 t^2}{2\ell_n^2} = \frac{n\nu_n t^2}{2} = \frac{t^2}{2\mu} + o(1). \quad (2.26)$$

Similarly, by the independence of $(T_j)_{j \in [n]}$,

$$\begin{aligned} \text{Var}(Y_0(t)) &= n^{-2/3} \text{Var}(N(sn^{2/3})) = n^{-2/3} \sum_{j \in [n]} \mathbb{P}(T_j \leq tn^{2/3})(1 - \mathbb{P}(T_j \leq tn^{2/3})) \\ &\leq n^{-2/3} \sum_{j \in [n]} \frac{w_j t n^{2/3}}{\ell_n} = t. \end{aligned} \quad (2.27)$$

Now we use the supermartingale inequality [25, Lemma 2.54.5], stating that, for any supermartingale $Y = (Y(s))_{s \geq 0}$, with $Y(0) = 0$,

$$\varepsilon \mathbb{P}(\sup_{s \leq t} |Y(s)| > 3\varepsilon) \leq 3\mathbb{E}(|Y(t)|) \leq 3 \left(|\mathbb{E}(Y(t))| + \sqrt{\text{Var}(Y(t))} \right). \quad (2.28)$$

Equation (2.28) shows that, for any large A ,

$$\mathbb{P}(\sup_{s \leq t} |N(sn^{2/3}) - sn^{2/3}| > 3An^{1/3}) \leq \frac{3(t^2/2\mu + t)}{A} + o(1). \quad (2.29)$$

This clearly proves that, for every $t > 0$,

$$\sup_{u \leq 2t} |n^{-2/3} N(un^{2/3}) - u| \xrightarrow{\mathbb{P}} 0. \quad (2.30)$$

Observe that (2.30) also immediately proves that, **whp**, $N(2tn^{2/3}) \geq tn^{2/3}$.

To deal with $\tilde{H}_n(\nu)$, we define

$$Y_1(u) = \tilde{H}_n(u) - \mu_3(n)u, \quad (2.31)$$

where we write

$$\mu_3(n) = \sum_{j \in [n]} \frac{w_j^3}{\ell_n} = \frac{\sigma_3}{\mu} + o(1), \quad (2.32)$$

and note that $Y_1(u)$ is a supermartingale. Indeed, writing \mathcal{F}_t to be the natural filtration of the above process, we have, for $s < t$ and letting $V_s = \{v : T_v < sn^{2/3}\}$,

$$\mathbb{E}(Y_1(t) | \mathcal{F}_s) = Y_1(s) + \frac{1}{n^{2/3}} \sum_{j \notin V_s} w_j^2 \left(1 - \exp\left(-\frac{n^{2/3}(t-s)w_j}{\ell_n}\right) \right) - \mu_3(n)(t-s). \quad (2.33)$$

Now using the inequality $1 - e^{-x} \leq x$ for $x \in [0, 1]$ we get that

$$\mathbb{E}(Y_1(t) | \mathcal{F}_s) \leq Y_1(s), \quad (2.34)$$

as required. Again we can easily compute, using condition (a), that

$$\begin{aligned} |\mathbb{E}[Y_1(t)]| &= -\mathbb{E}[Y_1(t)] = \mu_3(n)t - n^{-2/3} \sum_{i \in [n]} w_i^2 (1 - \exp(-tn^{2/3}w_i/\ell_n)) \\ &= n^{-2/3} \sum_{i \in [n]} w_i^2 \left(\exp(-tn^{2/3}w_i/\ell_n) - 1 + \frac{tn^{2/3}w_i}{\ell_n} \right) \\ &\leq n^{-2/3} \sum_{i \in [n]} w_i^2 \frac{(tn^{2/3}w_i)^2}{2\ell_n^2} \leq n^{2/3}t^2 \sum_{i \in [n]} \frac{w_i^4}{2\ell_n^2} = o(n^{2/3}n^{1/3}) \sum_{i \in [n]} \frac{w_i^3}{\ell_n^2} = o(1). \end{aligned} \quad (2.35)$$

By independence,

$$\begin{aligned} \text{Var}(Y_1(t)) &= n^{-4/3} \sum_{j \in [n]} w_j^4 (1 - \exp(-tn^{2/3}w_j/\ell_n)) \exp(-tn^{2/3}w_j/\ell_n) \\ &\leq n^{-2/3}t \sum_{j \in [n]} \frac{w_j^5}{\ell_n} = o(1) \sum_{j \in [n]} \frac{w_j^3}{\ell_n} = o(1). \end{aligned} \quad (2.36)$$

Therefore, (2.28) completes the proof of (2.16).

The proof of (2.17) is a little easier. We denote

$$\mathcal{V}_i = \{v(j)\}_{j=1}^i. \quad (2.37)$$

Then, we compute explicitly

$$\mathbb{E}[w_{v(i)}^2 | \mathcal{F}_{i-1}] = \sum_{j \in [n]} w_j^2 \mathbb{P}(v(i) = j | \mathcal{F}_{i-1}) = \frac{\sum_{j \notin \mathcal{V}_{i-1}} w_j^3}{\sum_{j \notin \mathcal{V}_{i-1}} w_j}. \quad (2.38)$$

Now, uniformly in $i \leq sn^{2/3}$, again using condition (a),

$$\sum_{j \notin \mathcal{V}_{i-1}} w_j = \sum_{j \in [n]} w_j + O((\max_{j \in [n]} w_j)i) = \ell_n + o(n) \quad (2.39)$$

for every $i \leq sn^{2/3}$. Similarly, again uniformly in $i \leq sn^{2/3}$, and using that $j \mapsto w_j$ is non-increasing,

$$\left| \sum_{j \notin \mathcal{V}_{i-1}} w_j^3 - \ell_n \sigma_3(n) \right| \leq \sum_{j=1}^{sn^{2/3}} w_j^3 = o(n). \quad (2.40)$$

Indeed, by observing that condition (b) is convergence of distribution of w_{V_n} to W , combined with condition (c), it is a standard fact that (for non-negative random variables, see [4, Theorem 3.6 on p. 31]) this implies that the third powers $w_{V_n}^3$ are uniformly integrable, which yields

$$\lim_{K \rightarrow \infty} \lim_{n \rightarrow \infty} \frac{1}{n} \sum_{j \in [n]} \mathbb{1}_{\{w_j > K\}} w_j^3 = 0. \quad (2.41)$$

Another standard result says that (2.41) is equivalent to $\mathbb{E}[\mathbb{1}_{A_n} w_{V_n}^3] \rightarrow 0$ for any sequence A_n of events with $\mathbb{P}(A_n) \rightarrow 0$, see [13, Theorem 10.3.5 on p. 355]. For $A_n = \{V_n \leq sn^{2/3}\}$, this gives, for each K ,

$$\frac{1}{n} \sum_{j=1}^{sn^{2/3}} w_j^3 \leq \frac{1}{n} \sum_{j=1}^{sn^{2/3}} \mathbb{1}_{\{w_j \leq K\}} w_j^3 + \frac{1}{n} \sum_{j \in [n]} \mathbb{1}_{\{w_j > K\}} w_j^3 \leq K^3 sn^{-1/3} + \frac{1}{n} \sum_{j \in [n]} \mathbb{1}_{\{w_j > K\}} w_j^3 = o(1), \quad (2.42)$$

when we first let $n \rightarrow \infty$, followed by $K \rightarrow \infty$.

We conclude that, uniformly for $i \leq sn^{2/3}$,

$$\mathbb{E}[w_{v(i)}^2 \mid \mathcal{F}_{i-1}] = \frac{\sigma_3}{\mu} + o_{\mathbb{P}}(1). \quad (2.43)$$

This proves (2.17). \square

To complete the proof of Theorem 2.1, we proceed to investigate $c(i)$. By construction, we have that, conditionally on \mathcal{V}_i ,

$$c(i) \stackrel{d}{=} \sum_{j \notin \mathcal{V}_i} I_{ij}, \quad (2.44)$$

where I_{ij} are (conditionally) independent indicators with

$$\mathbb{P}(I_{ij} = 1 \mid \mathcal{V}_i) = \frac{w_{v(i)} w_j (1 + tn^{-1/3})}{\ell_n}, \quad (2.45)$$

for all $j \notin \mathcal{V}_i$. Furthermore, when we condition on \mathcal{F}_{i-1} , we know \mathcal{V}_{i-1} , and we have that, for all $j \notin \mathcal{V}_{i-1}$,

$$\mathbb{P}(v(i) = j \mid \mathcal{F}_{i-1}) = \frac{w_j}{\sum_{s \notin \mathcal{V}_{i-1}} w_s}. \quad (2.46)$$

Since $\mathcal{V}_i = \mathcal{V}_{i-1} \cup \{v(i)\} = \mathcal{V}_{i-1} \cup \{j\}$ when $v(i) = j$, this is all we need to know to compute conditional expectations involving $c(i)$ given \mathcal{F}_{i-1} .

Now we start to prove (2.9), for which we note that

$$\begin{aligned} \mathbb{E}[c(i) \mid \mathcal{F}_{i-1}] &= \sum_{j \notin \mathcal{V}_{i-1}} \mathbb{P}(v(i) = j \mid \mathcal{F}_{i-1}) \mathbb{E}[c(i) \mid \mathcal{F}_{i-1}, v(i) = j] \\ &= \sum_{j \notin \mathcal{V}_{i-1}} \mathbb{P}(v(i) = j \mid \mathcal{F}_{i-1}) \sum_{l \notin \mathcal{V}_{i-1} \cup \{j\}} \frac{\tilde{w}_l w_l}{\ell_n}. \end{aligned} \quad (2.47)$$

Then we split

$$\begin{aligned}\mathbb{E}[c(i) - 1 \mid \mathcal{F}_{i-1}] &= \sum_{j \notin \mathcal{V}_{i-1}} \mathbb{P}(v(i) = j \mid \mathcal{F}_{i-1}) \tilde{w}_j - 1 - \sum_{j \notin \mathcal{V}_{i-1}} \mathbb{P}(v(i) = j \mid \mathcal{F}_{i-1}) \tilde{w}_j \sum_{l \in \mathcal{V}_{i-1} \cup \{j\}} \frac{w_l}{\ell_n} \\ &= \mathbb{E}[\tilde{w}_{v(i)} - 1 \mid \mathcal{F}_{i-1}] - \mathbb{E}[\tilde{w}_{v(i)} \mid \mathcal{F}_{i-1}] \sum_{s=1}^{i-1} \frac{w_{v(s)}}{\ell_n} - \mathbb{E}\left[\frac{w_{v(i)}^2(1 + tn^{-1/3})}{\ell_n} \mid \mathcal{F}_{i-1}\right].\end{aligned}\quad (2.48)$$

By condition (a), the last term is bounded by $O((\max_{j \in [n]} w_j)^2 / \ell_n) = o(n^{-1/3})$ uniformly in $i \in [n]$, and is therefore an error term. We continue to compute

$$\begin{aligned}\mathbb{E}[\tilde{w}_{v(i)} - 1 \mid \mathcal{F}_{i-1}] &= \sum_{j \notin \mathcal{V}_{i-1}} \frac{w_j^2(1 + tn^{-1/3})}{\sum_{s \notin \mathcal{V}_{i-1}} w_s} - 1 \\ &= \sum_{j \notin \mathcal{V}_{i-1}} \frac{w_j^2(1 + tn^{-1/3})}{\ell_n} - 1 + \sum_{j \notin \mathcal{V}_{i-1}} \frac{w_j^2(1 + tn^{-1/3})}{\ell_n \sum_{s \notin \mathcal{V}_{i-1}} w_s} \left(\sum_{s \in \mathcal{V}_{i-1}} w_s \right).\end{aligned}\quad (2.49)$$

The last term equals

$$\mathbb{E}\left[\frac{\tilde{w}_{v(i)}}{\ell_n} \mid \mathcal{F}_{i-1}\right] \sum_{s=1}^{i-1} w_{v(s)},\quad (2.50)$$

which equals the second term in (2.48), and thus these two contributions cancel in (2.48). This exact cancelation is in the spirit of the one discussed below (1.47). Therefore, writing $\tilde{v}_n = v_n(1 + tn^{-1/3})$, uniformly for all $i = o(n)$,

$$\begin{aligned}\mathbb{E}[c(i) - 1 \mid \mathcal{F}_{i-1}] &= \sum_{j \notin \mathcal{V}_{i-1}} \frac{w_j^2(1 + tn^{-1/3})}{\ell_n} - 1 + o(n^{-1/3}) \\ &= \sum_{j \in [n]} \frac{w_j^2(1 + tn^{-1/3})}{\ell_n} - 1 - \sum_{j \in \mathcal{V}_{i-1}} \frac{w_j^2(1 + tn^{-1/3})}{\ell_n} + o(n^{-1/3}) \\ &= (\tilde{v}_n - 1) - \sum_{s=1}^{i-1} \frac{w_{v(s)}^2(1 + tn^{-1/3})}{\ell_n} + o(n^{-1/3}) \\ &= (\tilde{v}_n - 1) - \sum_{s=1}^{i-1} \frac{w_{v(s)}^2}{\ell_n} + o(n^{-1/3}).\end{aligned}\quad (2.51)$$

As a result, we obtain that, uniformly for all $k = o(n)$,

$$A_n(k) = \sum_{i=1}^k \mathbb{E}[c(i) - 1 \mid \mathcal{F}_{i-1}] = k(\tilde{v}_n - 1) - \sum_{i=1}^k \sum_{s=1}^{i-1} \frac{w_{v(s)}^2}{\ell_n} + o(kn^{-1/3}),\quad (2.52)$$

where the sum of the error terms constitutes an error term uniform in $i \in [n]$ due to the *uniformity* in the bound in (2.51). Thus, again employing the uniformity in our error bounds,

$$\bar{A}_n(s) = ts - n^{-1/3} \sum_{i=1}^{sn^{2/3}} \sum_{l=1}^{i-1} \frac{w_{v(l)}^2}{\ell_n} + o(1).\quad (2.53)$$

By (2.16) in Lemma 2.2, we have that

$$\sup_{t \leq u} |n^{-2/3} \sum_{s=1}^{tn^{2/3}} w_{v(s)}^2 - \frac{\sigma_3}{\mu} t| \xrightarrow{\mathbb{P}} 0, \quad (2.54)$$

so that

$$\sup_{t \leq u} \left| \bar{A}_n(s) - ts + \frac{s^2 \sigma_3}{2\mu^2} \right| \xrightarrow{\mathbb{P}} 0. \quad (2.55)$$

This proves (2.9).

The proof of (2.10) is similar, and we start by noting that (2.51) gives that, uniformly for all $k = O(n^{2/3})$,

$$B_n(k) = \sum_{i=1}^k \mathbb{E}[c(i)^2 | \mathcal{F}_{i-1}] - \mathbb{E}[c(i) | \mathcal{F}_{i-1}]^2 = \sum_{i=1}^k \mathbb{E}[c(i)^2 - 1 | \mathcal{F}_{i-1}] + O(kn^{-1/3}). \quad (2.56)$$

Now, as above, we obtain that

$$\begin{aligned} \mathbb{E}[c(i)^2 | \mathcal{F}_{i-1}] &= \sum_{j \notin \mathcal{V}_{i-1}} \mathbb{P}(v(i) = j | \mathcal{F}_{i-1}) \sum_{\substack{s_1, s_2 \notin \mathcal{V}_{i-1} \cup \{j\} \\ s_1 \neq s_2}} \frac{\tilde{w}_j w_{s_1}}{\ell_n} \frac{\tilde{w}_j w_{s_2}}{\ell_n} \\ &\quad + \mathbb{E}[c(i) | \mathcal{F}_{i-1}]. \end{aligned} \quad (2.57)$$

From (2.51) we see that $\mathbb{E}[c(i) | \mathcal{F}_{i-1}] = 1 + o(1)$ uniformly for all $i = o(n)$, so that

$$n^{-2/3} B_n(n^{2/3}u) = n^{-2/3} \sum_{i=1}^{n^{2/3}u} \mathbb{E}[w_{v(i)}^2 | \mathcal{F}_{i-1}] + o(1) = \frac{\sigma_3}{\mu} u + o(1), \quad (2.58)$$

where the last equality follows from (2.17) in Lemma 2.2. The proofs of (2.9), (2.10) and (2.11) complete the proof of Theorem 2.1. \square

3 Verification of conditions (b)–(c): Proof of Corollary 1.2

3.1 Verification of conditions for i.i.d. weights

We now check conditions (b) and (c) for the case that $\mathbf{w} = (W_i)_{i \in [n]}$ where $(W_i)_{i \in [n]}$ are i.i.d. random variables with $\mathbb{E}[W^3] < \infty$. Condition (b) follows from the a.s. convergence of the empirical distribution function, while (1.12) in condition (c) holds a.s. by the strong law of large numbers. Equation (1.10) in condition (c) holds in probability by the central limit theorem (even with $o_{\mathbb{P}}(n^{-1/3})$ replaced with $O_{\mathbb{P}}(n^{-1/2})$). For the bound (1.11) in condition (c) we use the Marcinkiewicz-Zygmund law of large numbers (see e.g. [20, Theorem 4.23]) that says: If $0 < p < 2$ and X_i are i.i.d. with $\mathbb{E}|X_i|^p < \infty$ and $\mathbb{E}X_i = 0$, then $n^{-1/p} \sum_{i \in [n]} X_i \rightarrow 0$ a.s. In the case in the present paper, W_i are i.i.d. with $\mathbb{E}|W_i|^3 < \infty$. Taking $X_i = W_i^2 - \mathbb{E}[W^2]$, and applying the Marcinkiewicz-Zygmund law with $p = 3/2$, yields

$$\sum_{i \in [n]} W_i^2 - n\mathbb{E}[W^2] = o(n^{2/3}) \text{ a.s.}, \quad (3.1)$$

which proves (1.11).

3.2 Verification of conditions for weights as in (1.4)

Here we check conditions (b) and (c) for the case that $\mathbf{w} = (w_i)_{i \in [n]}$, where w_i is chosen as in (1.4). We make use of the fact that (1.3) implies that $1 - F(x) = o(x^{-3})$ as $x \rightarrow \infty$, which, in turn, implies that (see e.g., [14, (B.9)]), as $u \downarrow 0$,

$$[1 - F]^{-1}(u) = o(u^{-1/3}). \quad (3.2)$$

To verify condition (b), we note that by [16, (3.2)], w_{V_n} has distribution function

$$F_n(x) = \frac{1}{n} (\lfloor nF(x) \rfloor + 1) \wedge 1. \quad (3.3)$$

This converges to $F(x)$ for every $x \geq 0$, which proves that condition (b) holds. To verify condition (c), we note that since $i \mapsto [1 - F]^{-1}(i/n)$ is monotonically decreasing, for any $s > 0$, we have

$$\mathbb{E}[W^s] - \int_0^{1/n} [1 - F^{-1}(u)]^s du \leq \frac{1}{n} \sum_{i \in [n]} w_i^s \leq \mathbb{E}[W^s]. \quad (3.4)$$

Now, by (3.2), we have that, for $s = 1, 2, 3$,

$$\int_0^{1/n} [1 - F^{-1}(u)]^s du = o(n^{s/3-1}), \quad (3.5)$$

which proves all necessary bounds for condition (c) at once.

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