Binary black holes population and cosmology in new lights: Signature of PISN mass and formation channel in GWTC-3

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ABSTRACT

The mass, spin, and merger rate distribution of the binary black holes (BBHs) across cosmic redshifts provide a unique way to shed light on their formation channel. Along with the redshift dependence of the BBH merger rate, the mass distribution of BBHs can also exhibit redshift dependence due to different formation channels and dependence on the metallicity of the parent stars. We explore the redshift dependence of the BBH mass distribution jointly with the merger rate evolution from the third gravitational wave (GW) catalog GWTC-3 of the LIGO-Virgo-KAGRA collaboration. We study possible connections between peak-like features in the mass spectrum of BBHs and processes related to supernovae physics and time-delay distributions. We obtain a preference for short-time delays between star formation and BBH mergers. Using a power law form for the time delay distribution \((t_d^{\text{min}})^d\) we find \(d < -0.7\) credible at 90% interval. The mass distribution of the BBHs could be fitted with a power-law form with a redshift-dependent peak feature that can be linked to the pair instability supernova (PISN) mass scale \(M_{\text{PISN}}(Z^*)\) at a stellar metallicity \(Z^*\). For a fiducial value of the stellar metallicity \(Z^* = 10^{-4}\), we find the \(M_{\text{PISN}}(Z^*) = 44.4^{+7.9}_{-6.3}\) M\(_\odot\). This is in accordance with the theoretical prediction of the lower edge of the PISN mass scale and differs from previous analyses. Although we find a strong dependence of the PISN value on metallicity, the model that we explored is not strongly favored over those that do not account for metallicity as the Bayes factors are inconclusive. In the future with more data, evidence towards metallicity dependence of the PISN will have a significant impact on our understanding of stellar physics.

Key words: gravitational waves, black hole mergers, cosmology: miscellaneous

1 INTRODUCTION

Gravitational wave (GW) observations bring a wealth of information to a broad range of topics ranging from astrophysics, cosmology, and fundamental physics. The first GW detection (Abbott et al. 2016b) opened a new way of observing the Universe. The latest measurements from the LIGO-Virgo-KAGRA (LVK) (Gregory 2010; Aso et al. 2013; Acernese et al. 2014; Asa et al. 2015; Abbott et al. 2016a, 2018; Tse et al. 2019; Akutsu et al. 2020; Abbott et al. 2021a) have detected 90 compact objects which constitute of binary neutron stars (BNSs), neutron star binary black holes (NS-BHs), and binary black holes (BBHs) (Abbott et al. 2021b). The observed GW sources give a direct probe to infer the mass distribution of the compact objects across a range of cosmic redshifts. The recent measurement by the LVK collaboration exhibit that the mass distribution of the BBHs shows a power-law+Gaussian (PLG) distribution (Abbott et al. 2019b, 2021c,d). Along with the mass distribution, a power-law model of the BBHs merger rate is inferred from LVK analysis (Abbott et al. 2019a, 2021e). The mass distribution of GW sources and the merger rate provides a direct way to understand the formation channel of BBHs if an underlying physical model can be inferred from observations.

The currently used phenomenological PLG model of mass distribution does not consider redshift evolution. However, the mass distribution of the astrophysical BBHs is likely to exhibit a redshift dependence due to the dependence of the black hole masses on stellar properties, such as the stellar metallicity (Bethe & Brown 1998; Portegies Zwart & Yungelson 1998; Heger & Woosley 2002; Belczynski et al. 2002; Dominik et al. 2012; Dominik et al. 2015; Mapelli et al. 2017; Sperla & Mapelli 2017; Giacobbo et al. 2018; Toffano et al. 2019; Farmer et al. 2019a; Renzo et al. 2020; Baxter et al. 2021; Mehta et al. 2022). One of the inevitable ways the BBHs distribution can get a complex redshift dependence...
is through a time delay acting between the binary formation and the merger. (Mukherjee 2022). This delay causes the population of BBHs at a given redshift to encode information of astrophysical processes and channels that were present at a different cosmic epoch. The time delay contribution will produce a BBH population at a merger redshift $z$ (that is non-trivial to describe with simple models) which is composed by black holes formed at different cosmic times with possibly different astrophysical formation channels. The time delay distribution and the dependence of the astrophysical processes on cosmic time can lead to a non-trivial BBH merger mass spectrum.

The mass distribution of the BBHs can also play an important role in inferring the cosmic expansion history (Taylor et al. 2012; Farr et al. 2019; Mastrogiorgi et al. 2021a; You et al. 2021; Mancarella et al. 2021; Mukherjee 2022; Leyde et al. 2022; Ezquiaga & Holz 2022). As the masses of the GW sources are redshifted ($m^{\text{det}} = (1 + z)m$), one can expect to infer the mass redshift from the mass distribution of the GW sources, if the mass distribution of the BBHs can exhibit a universal property or at least a standardized behavior. We can break the mass-redshift degeneracy and infer the cosmic expansion history from dark standard sires without applying the cross-correlation technique (Oguri 2016; Mukherjee & Wandelt 2018; Mukherjee et al. 2020, 2021a; Bera et al. 2020; Scelfo et al. 2020; Mukherjee et al. 2021b; Cañas-Herrera et al. 2021; Scelfo et al. 2022; Cigarrán Díaz & Mukherjee 2022; Mukherjee et al. 2022) or statistical host identification technique (Schutz 1986; MacLeod & Hogan 2008; Del Pozzo 2012; Arabsalmani et al. 2013; Gray et al. 2020; Fishbach et al. 2019; Abbott et al. 2021g; Soares-Santos et al. 2019; Finke et al. 2021; Abbott et al. 2020b; Palmese et al. 2021). However, if BBHs mass distribution exhibits redshift dependence due to its intrinsic dependence on the delay time distribution, then cosmic redshifts cannot be accurately inferred (Mukherjee 2022; Ezquiaga & Holz 2022), and it can bias the results if the redshift dependence of the mass distribution is not considered. Exploring the merger rate distribution to explore cosmology from dark sires is also studied for the third generation GW detectors (Ding et al. 2019; Ye & Fishbach 2021; Leandro et al. 2022).

In this paper, we make a first joint estimation of the BBH merger rate evolution, mass distribution, and metallicity dependence parameters, by allowing for the redshift dependence of the BBH mass distribution and $H_0$. This measurement makes it possible to also infer the delay time distribution of the BBHs in a consistent framework along with the cosmological parameters. The paper is organized as follows, In Sec. 2.1 we discuss the redshift dependence of the mass distribution of the BBHs and its merger rate. In Sec. 3 and Sec. 4 we discuss the basic Bayesian framework used in this analysis, the results from the joint estimation and we compare these results with the results inferred by the LVK collaboration. Finally, we conclude the analysis of this work and future prospects in Sec. 5.

2 MODELLING REDSHIFT DEPENDENCE OF THE BINARY BLACK HOLE SOURCE POPULATION

In this analysis we use the following model to describe the distribution of BBHs in terms of their source frame masses $m_1, m_2$, and merger redshift $z_m$:

$$p(m_1, m_2, z_m | \Phi) = p(m_1, m_2 | z_m, \Phi_m, \Phi_d, \Phi_{\text{nuis}}) p(z_m | \Phi_d, \Phi_c).$$

where $\Phi = \{ \Phi_m, \Phi_c, \Phi_d, \Phi_{\text{nuis}} \}$ are a set of population parameters, merger redshift governing the mass model ($\Phi_m$), cosmology ($\Phi_c$), time delay ($\Phi_d$) and a set of nuisance parameters ($\Phi_{\text{nuis}}$).

In the following section we explain in detail each of those terms individually.

2.1 The redshift dependence of the mass distribution: Mixing of black holes

The distribution of BBHs observed by LVK spans a range of redshift and masses which is currently modeled using different phenomenological models out of which the PLG model fits the data the best (Abbott et al. 2019b, 2021e,c,d) but does not explore the redshift dependence of the BBH mass distribution. In this paper, we explore how an astrophysically motivated mass distribution of the BBHs originating due to the effect of the time delay distribution agrees with the GWTC-3 data. The observed mass distribution of the BBHs is driven by the underlying astrophysical properties of the parent stars of the individual black hole and the mixing of black holes formed in different redshifts, both of which lead to a redshift dependence of the observed BBH mass distribution.

The redshift dependence of the observed BBH mass distribution in the mixing of binary black hole model, that we consider in this analysis, is due to three effects, (i) metallicity dependence of the pair-instability supernovae (PISN) mass scale, (ii) redshift evolution of the stellar metallicity, (iii) distribution of delay times between the formation of the stars that will later become BBHs, and for them to merge with another black hole. We will briefly describe below all these aspects.

(i) According to the PISN process, the mass distribution of BBHs is expected to feature a mass gap due to the mass loss of heavy stars (Spera & Mapelli 2017; Farmer et al. 2019a; Renzo et al. 2020). The mass loss during the PISN sets the lower limit of the mass gap at around $M_{\text{PISN}} = 45 M_{\odot}$. However, $M_{\text{PISN}}$ is also closely related to the stellar metallicity. It was shown (Spera & Mapelli 2017; Farmer et al. 2019a; Renzo et al. 2020) that $M_{\text{PISN}}$ varies less than 10% for a variation of the stellar metallicity $Z$ from $10^{-5}$ to $3 \times 10^{-3}$ from a 1-D stellar evolution model Modules for Experiments in Stellar Astrophysics (MESA) (Paxton et al. 2011, 2019). The stars with higher metallicity have a larger mass loss due to stellar winds, which leads to a lower value of the PISN mass scale, than stars formed with lower metallicity. As a result, the position of the $M_{\text{PISN}}$ will vary. Given the simplicity of current modeling of stellar winds in 1-D codes such as MESA and the lack of independent observations to determine the PISN mass scale, the dependence of $M_{\text{PISN}}$ on stellar properties is still subject to large uncertainties.

(ii) The metallicity in the Universe varies with redshift and also with the individual galaxies. The global evolution of the
stellar metallicity (Belczynski et al. 2002; Dominik et al. 2012; Dominik et al. 2015; Mapelli et al. 2017; Giacobbo et al. 2018; Safarzadeh & Farr 2019; Toffano et al. 2019) indicates that the Universe at high redshift has poor stellar metallicity than at low redshift. As a result, the BHs formed at high redshift may have a higher PISN mass scale than the BHs formed at low.

(iii) Finally, the BHs which we observe using GWs, are not the individual BHs, but binaries. Though the formation of a black hole takes only a few Myrs, a black hole requires much more time to form a binary and merge. As a result, there is a non-zero delay time between the formation of a star and the merging of BHs. This delay time depends on the formation channels of the BBHs (O’Shaughnessy et al. 2010; Banerjee et al. 2010; Dominik et al. 2012; Dominik et al. 2015; Mandel & de Mink 2016; Lamberts et al. 2016; Cao et al. 2018; Eldridge et al. 2019; Vitale et al. 2019; du Buisson et al. 2020; Santoliquido et al. 2021). Moreover, the delay time is not a fixed number for all the BHs, but rather it follows a distribution that is expected to be a power law from simulations. For a flat in the log-space distribution of the separation of the BHs, the delay time distribution is going to be $t_d^{-1}$ with a minimum delay time from a few hundreds of Myrs to a few Gyrs, depending on the formation channels. The current constraints from GWTC-2 (Fishbach & Kalogera 2021) and the stochastic GW background (Mukherjee & Silk 2021) are weak. In the future, data-driven measurement is possible and the stochastic GW background (Mukherjee et al. 2021c). The delay time distribution is taken to be a simple power-law function $W(t_d|m; z_m)$ function brings a breaking point $t_m$ at the mass distribution, after which the mass distribution is suppressed depending on the form of the delay time distribution, the dependence of the PISN mass scale on stellar metallicity and the redshift evolution of the stellar metallicity. It is evaluated from the combination of different $M_{\text{PISN}}$ values which are governed by the minimum delay time $t_d^{\text{min}}$, metallicity evolution $\gamma Z$ and dependence of PISN mass scale on metallicity $\alpha Z$. The evolution of the $M_{\text{break}}$ for different choices of these parameters can be seen in Fig. 1. At a given $z_m$, the value of $M_{\text{break}}(z_m)$ is the minimum of $M_{\text{PISN}}$ over the formation redshifts included in $z(z_m,t_d^{\text{min}})<z_f<z(z_m,t_d^{\text{max}})$, namely

$$M_{\text{break}}(z_m) = \min_{t_d \in [t_d^{\text{min}}, t_d^{\text{max}}]} M_{\text{PISN}}(z_f(t_d, z_m)).$$

Assuming the dependence of the PISN mass scale on metallicity $M_{\text{PISN}}(Z)$ studied by Spera & Mapelli (2017); Farmer et al. (2019a); Renzo et al. (2020) can be reliably scaled using a parameter $\alpha Z$ (Mukherjee 2022), we can then model it with metallicity as:

$$M_{\text{PISN}}(Z) = M_{\text{PISN}}(Z_0) - \alpha Z \log_{10}(Z/Z_0),$$

and for a power-law redshift evolution of the stellar metallicity (as supported by the current observations (Mannucci et al. 2010; Sommariva et al. 2012; Krumholz & Dekel 2012; Dayal et al. 2013; Madau & Dickinson 2014)), we can write the redshift dependence of metallicity as $\log(Z) = \gamma Z + \zeta$. Consequently, the previous equation can be written as:

$$M_{\text{PISN}}(Z) = M_{\text{PISN}}(Z_0) - \alpha Z [\gamma Z + \zeta - \log_{10}(Z_0)],$$

where $\zeta = 10^{-2}$ is taken to be a constant to match the low redshift measurement of the stellar metallicity $Z(z = 0) \approx 10^{-2}$ and $Z_0 = 10^{-4}$. We select this $Z_0$ value as it is a mid-value inside the range for which PISN dependence has been explored in 1-D stellar mass model (Farmer et al. 2019a). In our analysis, we sample for the value of $M_{\text{PISN}}(Z_0)$.

The above expression is written in terms of the global evolution of stellar metallicity with redshift. However, at any particular redshift, there is going to be additional variation in the stellar metallicity depending on the property of the
host galaxy. So, we would expect variation in the parameter $\gamma_2$ on the property of the host galaxy. However, currently, due to the poor sky location error of the GW sources, we cannot identify the host galaxy and hence cannot model the parameter $\gamma_2$ as a function of the galaxy. As a result, there can be an additional variation in the $\gamma_2$ that cannot be well modeled currently. Similarly, the parameter $\alpha_2$ controls the dependence of the PISN mass scale on the metallicity (Farmer et al. 2019a). Hence we treat both these parameters as nuisance parameters in this analysis. We choose wide priors on $\alpha_2, \gamma_2$ which broadly include expected values from works cited above.

The connection of the observed BBH mass distribution with the PISN mass scale: We model the distribution of BBHs in terms of their source frame masses $m_1, m_2$ (with $m_1 \geq m_2$) and merger redshift $z_m$ of the binary as

$$p(m_1, m_2, z_m, \Phi_m, \Phi_d, \Phi_{nuis}) = p(m_1 | z_m, \Phi_m) p(m_2 | m_1, \Phi_m) S_1 S_2,$$

(8)

The masses in the detector frame (or redshifted masses) are given by:

$$m_{\text{det}} = (1 + z_m)m,$$

(9)

where $m$ are the masses in the source frame. To capture the mass distribution of BBHs that originate from the BBH mass distribution of Eq. 2, we consider $p(m_1 | z_m, \Phi_m)$ to be given by a power-law distribution superpositioned with the distribution of a Gaussian peak (Talbot & Thrane 2018; Abbott et al. 2021h,d,b):

$$p(m_1 | z_m, \Phi_m) = (1 - \lambda_0)P(m_1 | M_{\text{min}}, M_{\text{max}}, -\alpha) + \lambda_0G(m_1 | M_{\text{break}}(z_m), \sigma_g),$$

(10)

where $\Phi_m = \{M_{\text{min}}, M_{\text{max}}, \alpha, \lambda_0, M_{\text{break}}(z_m), \sigma_g\}$, $G(m_1 | M_{\text{break}}(z_m), \sigma_g)$ is a Gaussian distribution with $\mu = M_{\text{break}}(z_m)$ and $\sigma = \sigma_g$ and $P(m_1 | M_{\text{min}}, M_{\text{max}}, -\alpha)$ is a power-law distribution with slope $-\alpha$ between $M_{\text{min}}$ and $M_{\text{max}}$. In this model the power-law part of the mass distribution is motivated by the power-law form of the initial mass function (IMF) Kroupa (2002) and the Gaussian part of the mass distribution is motivated by the PISN mass scale. The sources merging at redshift $z_m$ due to the contribution from all the higher redshift will lead to an excess near the value of $M_{\text{break}}$ and then a decline in the mass distribution due to the window function. The position of the Gaussian peak $\mu$ is considered at the break of the window function at that redshift, which depends on the metallicity dependence of the PISN mass scale and delays time distribution. The value of the PISN mass scale is inferred for the metallicity value at $Z = 10^{-4}$ (for which the results are obtained by Farmer et al. (2019b)). The Gaussian peak modeled in this analysis gets a physical motivation expected from the PISN mass scale but is also expected to evolve as a function of the redshift of BBHs mergers.

The distribution of $m_2$ in the source frame is considered to be given by a power law with maximum value $m_1:

$$p(m_2 | \Phi_{m_2}) = P(m_2 | M_{\text{min}}, m_1, \beta).$$

(11)

Since $m_2$ is conditional to $m_1$, the window function $W_{td}$ is being applied also to $m_2$ indirectly.

Finally, the functions $S_{(1,2)} = S(m_{(1,2)} | \delta_m, M_{\text{min}})$ are sigmoid-like window functions to smooth the lower end of the distributions (see appendix A of (Abbott et al. 2021d)). We choose to consider only the position of the Gaussian peak to vary with redshift since this is the most prominent feature in the mass spectrum of BBHs and is the best-constrained parameter. Other mass parameters of the mass model $(M_{\text{max}}, \alpha, M_{\text{min}}, ...)$ can also be given a redshift or metallicity dependence (van Son et al. 2022). However currently, with the limited number of GW sources, measurement of the redshift dependence of the additional parameters will be difficult or unlikely to be strongly constraining.

2.2 The redshift dependence of the BBH merger rate distribution

The distribution $p(z_m | \Phi_d, \Phi_c)$ takes into account the BBH merger rate as a function of redshift and it is built as

$$p(z_m | \Phi_d, \Phi_c) = C \frac{R(z_m) dV_c}{1 + z d^2dz_m},$$

(12)

where $C$ is a normalization constant, $\frac{dV_c}{dz_m}$ the differential of the comoving volume and $R(z_m)$ the BBH merger rate as function of redshift. The BBH merger rate is built as

$$R(z_m) = R_0 \frac{\int_{z_m}^{\infty} P(t_d | t_{d_{\text{min}}}, t_{d_{\text{max}}}, d) R_{\text{SFR}}(z_f) \frac{dz_f}{dz_m} dz_f}{\int_{0}^{\infty} P(t_d | t_{d_{\text{min}}}, t_{d_{\text{max}}}, d) R_{\text{SFR}}(z_f) \frac{dz_f}{dz_m} dz_f},$$

(13)

where $z_f$ is the redshift of formation, $R_0$ the BBH merger rate today, $P(t_d | t_{d_{\text{min}}}, t_{d_{\text{max}}}, d)$ is a time-delay distribution between formation and merger of the binary and $R_{\text{SFR}}(z_f)$ is a parametrization for the Madau-Dickinson star formation rate (SFR) (Madau & Dickinson 2014).

We show the BBH merger rate for a few values of the variables $d$, $t_{d_{\text{min}}}$ and a fiducial flat Lambda Cold Dark Matter (LCDM) cosmology model with $H_0 = 70$ km/s/Mpc and $\Omega_m = 0.3$ in Fig. 2 with $R_0 = 20$ Gpc$^{-3}$ yr$^{-1}$. The plot indicates that with the decrease in the value of power-law index $d$ and minimum delay time $t_{d_{\text{min}}}$, the peak in the BBH merger shifts towards a higher redshift with a steeper slope at low redshift. Current observations from LVK can measure BBHs at redshifts ($z < 1$) for a fiducial model of cosmology.

The above discussion shows that the delay time distribution $P(t_d | t_{d_{\text{min}}}, t_{d_{\text{max}}}, d)$ plays a role in both the mass distribution of the BHs and also in their merger rates. As a result, to infer the BBH formation channel and the delay time distribution, it will be necessary to use both merger rate and mass distribution to infer the delay time distribution of BBHs. Moreover, to estimate the redshift of the BHs from their mass distribution, one needs to account for the redshift dependence of the BBH mass distribution. As a result, we need to jointly infer the cosmological parameters along with the delay time distribution of BHs, and the black hole mass distribution, to correctly marginalize the degenerate parameters.

3 BAYESIAN FRAMEWORK TO INFER COSMOLOGY FROM GW POPULATION

We construct a Bayesian model following the method described in (Mastrogiavanni et al. 2021b; Abbott et al. 2021d)
Fixing $d = -1$ and varying $t_{d}^{\text{min}}$. On the plots, the star formation rate $R_{\text{SFR}}/R_{0,\text{SFR}}$ can also be seen.

Figure 2. The merger rate function $R(z)/R_0$ for various values of the parameters $d$, $t_{d}^{\text{min}}$ and fiducial flat Cold Dark Matter cosmology with a constant energy density for dark energy, $H_0 = 70 \text{ km/s/Mpc}$ and $\Omega_m = 0.3$. Left: Fixing $t_{d}^{\text{min}} = 0.5$ Gyrs and varying $d$. Right: Fixing $d = -1$ and varying $t_{d}^{\text{min}}$. On the plots, the star formation rate $R_{\text{SFR}}/R_{0,\text{SFR}}$ can also be seen.

Figure 3. Flow chart of the inference method.

to conjointly infer the redshift dependence of the mass distribution and merger rate along with the cosmological parameters. In Fig. 3 we show a schematic diagram explaining the formalism. As a cosmological model, we consider a flat LCDM (Aghanim et al. 2018). Given a set of $N$ GW detections associated with the data $\{x\} = \{x_1, \ldots, x_N\}$, the posterior on $\Phi$ can be expressed as (Thrane & Talbot 2019; Mandel et al. 2019; Mastrogiavanni et al. 2021b; Abbott et al. 2021d; Vitale et al. 2021)

$$p(\Phi, \{x\}, N) = \Pi(\Phi) e^{-N_{\text{exp}}(\Phi)} N_{\text{exp}}(\Phi) N_{\text{obs}} \times \prod_{i=1}^{N} \int p(x_i|\Phi, \theta) p_{\text{pop}}(\theta|\Phi) d\theta$$

where $\Pi(\Phi)$ is prior on the parameters, $\theta$ is the set of intrinsic parameters, which are unique for each event, $p(x_i|\Phi, \theta)$ is the likelihood, $p_{\text{det}}(\theta, \Phi)$ is the probability of detection and $p_{\text{pop}}(\theta|\Phi)$ is the population modeled prior. Finally, $N_{\text{exp}}(\Phi)$ is the expected number of detections in a given observing time, and $N_{\text{obs}}$ is the number of events considered in the analysis. The term $p_{\text{pop}}(\theta|\Phi)$ is given by Eq. 1. This term captures the effects of delay time between formation and merger.

The denominator of Eq. 14 normalizes the numerator and takes into account the selection effects (Mastrogiavanni et al. 2021b; Abbott et al. 2021d). It is written as an integral over all detectors’ data that pass certain detection criteria for given known noise properties of the detectors. The term $p_{\text{det}}(\theta, \Phi)$ is the probability of detecting the source with parameters $\theta$ and assuming hyper-parameters $\Phi$.

The summary of the priors used for the parameters that we consider in our model can be found in Table 1.

4 RESULTS FROM GWTC-3

We analyzed all the BBH events from GWTC-3 (Abbott et al. 2021e) with the matched filtering Signal-to-Noise Ratio (SNR) in at least one of the detection pipelines higher than 12. We select a high SNR threshold to avoid any possible contamination from noise. Moreover, the choice of a high SNR threshold is motivated by the fact that our selection biases are evaluated with injections in simulated and not real data. Detection properties between simulated data and real data might be different, especially when lowering the threshold for detection. For all the events, we also require a False Alarm Rate $< 0.25\text{ yr}^{-1}$.

The injection campaign is done in simulated data with duration and sensitivity typical of O1, O2, and O3. For all the observing runs, we assume independent duty cycles among the LIGO Hanford (H1), LIGO Livingston (L1), and Virgo (V1) detectors taken from (Abbott et al. 2021f) for O1 and O2 and O3 (Buikema et al. 2020; The LIGO Scientific Collaboration et al. 2023). For O1 we assume duty cycles of 64.6% for H1 and 57.4% for L1, while in O2 it was 65.3% and 61.8%. For the entire O3, we assumed 74.6% for H1, 77.0% for L1, and 76.0% for V1. We used the power spectral density of the publicly available detectors for O1 $^1$ and O2 $^2$, while for O3 we used an estimation provided with the first three months of O3a $^3$. Moreover we have assumed the noise of the detector to be Gaussian and stationary. The injections are performed using the IMRPHENOMPv2 waveform model and are drawn from a distribution in detector frame masses and luminosity distance large enough to cover all the detectable sources.

$^1$ https://www.gw-openscience.org/O1/
$^2$ https://www.gw-openscience.org/O2/
$^3$ https://dcc.ligo.org/LIGO-T2000012/public

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### Delay time + Merger rate parameters

| Parameter | Description | Prior |
|-----------|-------------|-------|
| $d$       | Spectral index for the power-law of the delay time distribution. | $U(-4,0)$ |
| $\mu_{\text{min}}^d$ | Minimum time for the power-law of the delay time distribution in Gyrs. | $U(0.01,13)$ |
| $R_0$     | Value of the merger rate at $z = 0$ in $\text{Gpc}^{-3}\text{yr}^{-1}$. | $U(0,1000)$ |

### Mass distribution parameters

| Parameter | Description | Prior |
|-----------|-------------|-------|
| $\alpha$  | Spectral index for the power-law of the primary mass distribution. | $U(-4,12)$ |
| $\beta$   | Spectral index for the power-law of the mass ratio distribution. | $U(-4,12)$ |
| $M_{\text{min}}$ | Minimum mass of the power-law component of the primary mass distribution in $M_\odot$. | $U(20,10)$ |
| $M_{\text{max}}$ | Maximum mass of the power-law component of the primary mass distribution in $M_\odot$. | $U(50,200)$ |
| $\lambda_9$ | Fraction of the model in the Gaussian component. | $U(0,1)$ |
| $M_{\text{PISN}}(Z_*)$ | The value of $M_{\text{PISN}}$ for the metallicity value $Z_*$ in $M_\odot$. | $U(20,60)$ |
| $\sigma_8$ | Width of the Gaussian component in the primary mass distribution in $M_\odot$. | $U(4,10)$ |
| $\delta_m$ | Range of mass tapering at the lower end of the mass distribution in $M_\odot$. | $U(0,10)$ |

### Cosmological parameters (Flat LCDM model)

| Parameter | Description | Prior |
|-----------|-------------|-------|
| $H_0$     | The Hubble constant parameter in km/s/Mpc. | 67.4(fixed); $U(20,150)$ |
| $\Omega_m$ | Present-day matter density of the Universe. | 0.315 (fixed) |

### Nuisance parameters

| Parameter | Description | Prior |
|-----------|-------------|-------|
| $\alpha_Z$ | The parameter that captures a weak logarithmic dependence of $M_{\text{PISN}}$ on the metallicity. | $U(-15,15)$ |
| $\gamma_Z$ | The parameter that captures the redshift dependence of the metallicity. | $U(-5,0)$ |
| $\zeta$   | The parameter that captures the metallicity at redshift $z = 0$. | 0.01 (fixed) |

Table 1. Summary of the hyperparameters and the priors used. The distribution $U(\text{min}, \text{max})$ is just a uniform distribution between min and max for each parameter. Note that the breaking mass $M_{\text{break}}(z_m)$ is implied by the choice of the other population parameters.

Influence the inferred values of the GW source parameters, given the current precision of the cosmological parameters from Planck-2018, the expected statistical error in the inferred GW source parameters is much larger than the systematic error due to different choices in the value of $H_0 = 66.9$ km/s/Mpc (Aghanim et al. 2018) or $H_0 = 73$ km/s/Mpc (Riess et al. 2021). Case-3: as Case-2 but keeping the value of the delay time power-law index fixed and equal to $d = -1$, which is usually assumed as a fiducial scenario for flat in the log-space distribution of the separation between the binaries.

The analysis for Case-3 (with a fixed value of $d = -1$) is motivated to find the constraints on the parameter space for the fiducial scenario of the delay time distribution, which resembles closely previous analysis (Abbott et al. 2021b). The priors of the runs for each of the parameters can be found in Table 1. We have summarised the estimated values of the parameters for different cases in Table 2. All the quoted values of the error bars are 68% confidence intervals unless mentioned otherwise.

**Case-1 (GW source population + $H_0$):** For this scenario, the joint constraints on the thirteen GW source population parameters and one cosmological parameter, $H_0$, is shown in Fig. 5. The joint estimation can be broadly classified into the parameters related to the delay time + merger rate, mass distribution, and cosmology. Among the delay time + merger rate parameters, we find that data support a scenario of a steep increase in the redshift evolution of the merger rate ($d < -1$), with BBHs merger rate density at $z = 0$ $R_0 = 22.3^{+7.5}_{-6.7}$ Gpc$^{-3}$ yr$^{-1}$. Our constraint on the GW merger rate at high redshifts is dominated by our assumptions on the SFR and the constraints we obtain on the time-delay parameters at low redshifts. However, if there is a different population of BBHs that do not contribute to low redshifts according to the Madau-Dickinson SFR, but contributes to the high redshift such as the Pop-III star, that cannot be constrain from this analysis. The expected number of events after including the detector noise and duty cycle agrees well with the total number of events with SNR $\geq 12$ considered in this analysis.

Using this model, we impose constraints on the minimum delay time distribution $t_d^{\text{min}} < 10$ Gyrs$^4$. We find the power-law index of the delay time distribution to be constrained $d < -0.7$ and there is a mild preference towards values lower than $d = -1$ as expected from a simple scenario of flat in the log-space distribution of object separation of the binaries. This measurement shows a steep distribution of the delay time and hints towards scenario formation channels having

4 Upper or lower limits are based on the 90% CL.
measurements around the median are based on the 68% CL.

The constraints on the $t_{\text{d}}^\text{min}$ and $d$ obtained here are driven by the joint estimation of the merger rate and mass distribution of the BBHs. Larger (smaller) values of $t_{\text{d}}^\text{min}$ or larger (smaller) values of $d$ support higher (lower) delay time values in $P(t_{\text{d}})$. Mergers of BBH from higher redshifts are supported from the scenarios with large values of $t_{\text{d}}^\text{min}$ or large values of $d$. The BBHs with heavier component masses in the data are fitted with BHs appearing from a higher redshift with higher PSN masses and a non-zero value of the delay time.

In our analysis, we also constrain models with a metallicity evolution in the Universe through the parameter $\gamma_Z$. The value of the parameter $-\gamma_Z = 3.2^{+1.1}_{-1.2}$ shows that there is likely an evolution of the metallicity of the parent stars. In comparison to the current observations, (Mannucci et al. 2010; Sommariva et al. 2012; Krumholz & Dekel 2012; Dayal et al. 2013; Madau & Dickinson 2014) and also proposed from simulations (Genel 2016; Torrey et al. 2019), the posterior on $\gamma_Z$ is consistent, supporting a decrease in the stellar metallicity with redshift. However, depending on the metallicity of the host galaxy, the parameter $\gamma_Z$ can have additional dependence on the astrophysical property of the host galaxy (Artale et al. 2019, 2020), and hence can exhibit additional variation from the mean metallicity value of the Universe. Such effects can show up when more events are available and hence better modeling of the BBHs population will be needed.

The parameters related to the mass distribution are also constrained well using a model. We have obtained a value of the power-law index of the mass distribution $\alpha = 3.2^{+1.0}_{-0.8}$.
and $\beta = 1.0^{+1.2}_{-0.9}$. We also find support for a feature over a simple power-law in the mass spectrum of BBH mergers with a relative height of the feature with respect to the power-law component of $\lambda_g < 0.13$ and the position of the feature is inherited at a fiducial metallicity from $M_{\text{PISN}}(Z_*) = 46.8^{+6.8}_{-7.3} M_\odot$ at $Z_* = 10^{-4}$. Differently from (Abbott et al. 2021b), we are not able to exclude with confidence the value $\lambda_g = 0$ (absence of a peak feature). This is due to the use of selection criteria based on a higher SNR cut instead of an IFAR cut. We verified that with a vanilla PLP model and an IFAR cut of 1 yr as in (Abbott et al. 2021b), we can exclude the absence of the peak.)

The position of the peak agrees with the theoretically predicted position of the PISN mass scale between $45$-$60$ $M_\odot$ (Farmer et al. 2019a; Renzo et al. 2020; Baxter et al. 2021). In Fig. 9, we show the PISN position is translated to the BBH merger primary mass spectrum when taking into account the full-time-delay model. As we can see from the figure, the PISN mass scale between $45$-$60$ $M_\odot$ is translated to a BBH merger excess at around $35$ $M_\odot$ for $z < 0.2$. This is compatible with the overdensity of BBHs observed by the

Figure 5. Posterior distributions for all the hyperparameters. Here we have fixed all cosmological parameters besides $H_0$ in Planck-2018 cosmology. We have used all GW events with $SNR \geq 12$. This plot corresponds to case 1 mentioned in the results section.
Figure 6. Posterior distributions for all the hyperparameters. Here we have fixed all cosmological parameters at Planck-2018 cosmology.
We have used all GW events with $SNR \geq 12$. This plot corresponds to case 2 mentioned in the results section.

LVK in Abbott et al. (2021b). However, as redshift increases this moves to higher masses and appears to become more prominent.

For high-mass BBH, we see a significant increase in merger rate with redshift, clearly indicating that there is support for a higher merger rate from high masses at higher redshift in comparison to the low redshift. The posterior distribution of the $M_{\text{PISN}}$ parameter is shown separately in Fig. 11 (bottom, blue curve). The PISN mass scale depends on the value of metallicity and this value of $M_{\text{PISN}}$ is defined in our analysis at the value of $Z_\ast = 10^{-4}$, which is in agreement with

the parameters chosen in the simulation by (Farmer et al. 2019a). The dependence of the PISN mass scale on metallicity is stronger, i.e. the probable values of $\alpha Z$ are larger, in this model than is expected from the 1-D stellar evolution models of Paxton et al. (2011, 2019).

Also, as it is evident from Fig. 5 the data has strongly suppressed any negative values of $\alpha Z$. So, scenarios of a decrease in the PISN mass scale with a decrease in the metallicity are strongly ruled out.

It is important to note that the current theoretical estimation on the dependence of the PISN mass scale is sub-
However, the data were consistent with both an evolving and a truncated power law with a sharp cut-off at high masses.

For the redshift evolution of the mass model when considering GWTC-2 (Fishbach et al. 2021). They found strong evidence for the redshift evolution of the mass model when considering a first-generation BBH formation scenario, one needs a stronger dependence of the PISN mass scale on the stellar metallicity. This can be seen in Fig. 7. We find that most of the parameters do not show up significant deviation from previous results (Abbott et al. 2021b,d). However, differently from (Abbott et al. 2021b,d), $\sigma_d$ is not well constrained. The high $\sigma_d$ estimation is an indication that there may not be a Gaussian feature in the mass distribution and the mass distribution can be smeared with an extended distribution in the masses at the higher end.

Finally, a weak measurement of the Hubble constant $H_0 = 42^{+19}_{-12} \text{ km/s/Mpc}$ is made which is in the agreement with the values from Planck-2018 (Aghanim et al. 2018), SH0ES (Riess et al. 2021) due to large uncertainty in the current measurements. However, note the correlation between the Hubble constant and the $t_d^{\text{min}}$ parameter in Fig. 5. This indicates that not being able to correctly infer the PISN mass scale and the value of $t_d^{\text{min}}$ can lead to an incorrect inference of the cosmological parameters (Mukherjee 2022).

**Case-2 (Main model):** In this case, we only focus on the GW source population keeping the value of cosmological parameters fixed at the Planck-2018 (Aghanim et al. 2018). The corresponding joint estimations of the parameters are shown in Fig. 6. From the posteriors, we can obtain the merger rate model for various samples, along with the median. This can be seen in Fig. 7. We find that the value of the $M_{\text{PISN}}(Z_d) = 44.4^{+7.2}_{-6.3} M_\odot$, has moved to lower values with respect to the value allowing $H_0$ to vary. The value of $t_d^{\text{min}}$ shows a maximum a posteriori around 1.5 Gyrs and a significantly narrower posterior with respect to Case-1. The posteriors for $r_d^{\text{min}}$ and $d$ are consistent with the case of varying $H_0$. We also find $d < -0.87$ for this case. We retrieve weak evidence of $M_{\text{PISN}}(Z_d)$ evolving with redshift as the $\alpha_d$ parameter supports positive values. The value of $M_{\text{PISN}}(Z_d)$ spans from around 30 M$_\odot$ for $z = 0$ up to around 40 M$_\odot$ for $z = 1$. As shown in Fig. 8, the redshift evolution of the PISN mass scale shows a weak variation over the redshift range $z \in [0, 1]$. However, more observation will be required to confidently make any detection of the redshift evolution of PISN mass distribution. A few samples in Fig. 8 show $M_{\text{PISN}}(z)$ values around 80 M$_\odot$ at redshift $z = 0$ and exhibit a decay to a non-evolving mass distribution when they considered a broken power-law model as a mass model. Those findings are broadly in agreement with our results.

We find that most of the parameters do not show up significant deviation from previous results (Abbott et al. 2021b,d). However, differently from (Abbott et al. 2021b,d), $\sigma_d$ is not well constrained. The high $\sigma_d$ estimation is an indication that there may not be a Gaussian feature in the mass distribution and the mass distribution can be smeared with an extended distribution in the masses at the higher end.
Figure 10. Posterior distributions for all the hyperparameters. Here we have fixed all cosmological parameters at Planck-2018 cosmology. We have fixed the power-law index of the delay time to $d = -1$. We have used all GW events with $SNR \geq 12$. This plot corresponds to case 3 mentioned in the results section.

crease in the $M_{\text{PISN}}(z)$ with redshift evolution. Those arise from the tail of the posterior distribution due to statistical fluctuations.

Comparison of the simple PLG peak model with our model we retrieve a Bayes factor (BF$_{\text{PLG/main model}}$) equal to $\log_{10}(\text{BF}_{\text{PLG/main model}}) = 0.32$ in favor of the simple PLG peak model. So, we conclude that there is a slight but insignificant preference for the PLG model. A higher number of detections is required to obtain decisive evidence for or against the redshift evolution of the BBH mass distribution.

We also compare the results with the events from GWTC-3 with $SNR \geq 11$ in Fig. A1 (shown in the appendix). The results are consistent with the measurement of the parameters made with events having an $SNR \geq 12$. In this analysis, we have not considered the GW event GW190521 (Abbott et al. 2020a) which has a much higher value of the component masses. Results including GW190521 can be seen in the appendix (see in Fig. B1). Constraints on the GW source parameters are very similar for both with or without GW190521.

To explore whether the model struggles to fit high masses seen in the data, we also consider a variation of our main
model. We refer to this modification as the “High mass” model and in this we impose the window function $W_{t_d}$ only in the Gaussian peak of the distribution, leaving the power-law intact. For a given redshift, we can fit higher mass events but also consider a fiducial fixed delay time power-law index $d = -1$ and estimate the rest of the population parameters (labeled as Pop). Case 3: Keep all the cosmological parameters fixed to Planck 2018 values but also consider a fiducial fixed delay time power-law index $d = -1$ and estimate the rest of the population parameters (labeled as Pop(d=-1)).

Case-3 (GW source population (with fixed $d = -1$) parameters): In Case-1 and Case-2, we have seen a value of the power-law index significantly away from $d = -1$ (usually considered as a fiducial value (Fishbach & Kalogera 2021; Mukherjee & Silk 2021)), which is possible for scenarios with flat in log space distribution of the separation between the BHs. Here we perform a joint estimation with the value of the power-law index fixed at $d = -1$. The corresponding joint estimations of the parameters are shown in Fig. 10. For this case, the merger rate normalization is $R_0 = 27.3^{+7.4}_{-5.5}$ Gpc$^{-3}$ yr$^{-1}$, and the constraints on the minimum delay time $t_d^{\text{min}}$ have been reduced (though completely in agreement with the value obtained with $d$ varying from the two previous cases). This happens because, for a scenario with a fixed value of the parameter $d = -1$, the peak of the merger rate distribution is shifted towards a lower redshift, as a result by allowing a smaller value of the delay time $t_d^{\text{min}}$, the peak of the merger rate position shifts towards a higher redshift. The position of the Gaussian peak has a very similar value $M_{\text{PSN}}(Z_*) = 43.8^{+7.5}_{-6.5}$ M$_\odot$ with respect to the previous results of Case-2.

Among all our time delay models, we find that the preferred model is Case-2, namely the case in which cosmology is fixed but the index of the time delay distribution $d$ varies. More interestingly, the Case-3 model ($d = -1$) is disfavored with respect to Case-2 by a Bayes factor of $\log_{10}(BF_{\text{vary } d/\text{fixed } d}) = 0.38$. However, this is not enough to claim any statistical evidence and we can not conclude any preference towards against or in favor of $d = -1$. The posteriors of $t_d^{\text{min}}$ and $M_{\text{PSN}}(Z_*)$ can be seen in the top and bottom panels of Fig. 11 for the three main cases that we considered.

The results obtained using our model including the delay time distribution and redshift evolution in the merger rate indicate a value of $M_{\text{PSN}} = 44.4^{+7.9}_{-6.3}$. In this model, the heavy mass BHs are formed from the low metallicity parent stars at a high redshift that has merged at a low redshift due to a delay time distribution function which allows large values of delay time. The Bayes factor in favor of this model in comparison to the phenomenological model of PLG is comparable and cannot be well distinguished at this stage. However, in the future with more data, we will likely be able to distinguish between different scenarios.

One of the major drawbacks of the baseline model considered here is that it only considers first-generation BBHs and not the scenarios where the second-generation BBHs are present. As a result, the presence of heavier masses observed in GWTC-3 is expected to arise from stars with low metallicity. However, in reality, there can be second-generation BBHs that can contribute to the observed population. A successful physics-driven model needs to consider this aspect as well. We will explore this in a future work.

5 CONCLUSION

The mass, spin, and merger rate of BBHs are a direct probe to infer their formation channels. Though the information available from observations of BBH component spins is limited (see e.g. Abbott et al. (2021b)), we can infer the masses and luminosity distance of several BBHs using the network of LIGO/Virgo detectors. The mass distribution and merger rate of the BBHs are likely to exhibit a redshift dependence due to the dependence on the stellar metallicity and forma-
tion channel of BBHs. In this work, we consider a model which considers the redshift dependence of BBH mass distribution due to the mixing of binary black holes and the redshift dependence of merger rate. We used a Bayesian analysis and estimated the values of the model’s parameters using the latest GW catalog of LVK GWTC-3.

By fitting this model with the data with both GW source parameters and cosmological parameters. We show that using a time-delay distribution for BBHs and a PISN mass model in redshift, it is possible to obtain a value of $M_{\text{PISN}}(Z_\odot)$ compatible with astrophysical expectations. This shows that the BBH excess in the mass profile of BHs could likely be reconciled with the PISN mass-scale value of around $40 - 50 M_\odot$ when time delay information is considered. However, if there exist BHs of second-generation that are going to have masses heavier than predicted by the PISN mass scale, then such systems cannot be captured by this model. Though we have found that current data equally favors a model allowing only for first-generation BBH mergers with redshift-dependent mass distribution over the phenomenological redshift-independent PLG mass model. There also is an indication of the redshift dependence of the BBH mass distribution, but this trend is not statistically significant. More observations will be required to better understand the redshift dependence. Our analysis also indicates a value of $\alpha Z = 7.0_{-2.9}^{+4.0}$ for the Case-2 (fixed $H_0$ case). This variation is $3 - 4$ times larger than the typical value people expect from 1-D stellar simulations and a simplistic prescription of stellar wind models Farmer et al. (2019b). In the future with more data, if these results hold, then one needs to better understand the dependence of metallicity on $M_{\text{PISN}}(Z_\odot)$.

In the hypothesis that the BBHs we consider are formed in a stellar binary scenario, we provide an upper limit for the minimum of the time delay distribution. The delay time distribution agrees with the formation channels explored in the literature. We find support for values of the power-law index of the delay time distribution $d < -1$. The value $d = -1$ is usually considered as the fiducial value for flat in-log space distribution of the spacing between the binaries. We retrieve values $d < -1$ for the power index of the time delay distribution. Though we can not exclude with certainty the fiducial scenario $d = -1$, we do find a mild preference for smaller values.

In our analysis, we also jointly infer the value of the Hubble constant, which currently exhibits a large uncertainty and hence is consistent with the value of the Hubble constant inferred from Planck-2018 and SHOES. However, one of the important parts is that the Hubble constant shows strong degeneracy with the GW source parameters. In particular, in this paper, we have shown that the estimation of the Hubble constant could also be impacted by the BBHs time delay distribution. As a result, if the inferred value of the GW source population is incorrect, then it can bias the value of the Hubble constant and other cosmological parameter estimation.

In the future with the availability of more sources from the next observation run, a better measurement of the GW source population along with the cosmological parameters will be possible. This will shed light on the formation channels of the binary systems and the redshift dependence of the BH mass distribution. Also, improvement in the theoretical modeling will be required to capture the underlying distribution of the BBHs from both the first generation and second generations.

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DATA AVAILABILITY

The data underlying this article is available on this website https://www.gw-openscience.org.

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APPENDIX A: PARAMETER ESTIMATION WITH A DIFFERENT SNR THRESHOLD.

Here we explore the variation of our results in the case of a lower SNR threshold. In addition to the GW events that we used with SNR $\geq 12$ (see Sec. 4), we are now using additionally all events with SNR $\geq 11$. Lowering the SNR causes to run the analysis with 41 events in total. The different posteriors can be seen in Fig. A1. It is apparent that all of the posteriors are in agreement with the ones obtained excluding GW190521. We find a slightly less steep power-law for the $m_1$ distribution with respect to the run excluding GW190521 with a power-law index for the mass distribution of $\alpha = 3.0_{-0.7}^{+0.9}$. We also retrieve a flatter $\sigma_g$ posterior and a posterior for $M_{\text{max}}$ that disfavors smaller values of masses, though fully in agreement with the previous results. These differences are because GW190521 components are very massive. For the model to fit the extra support at high masses, it needs a less steep power-law index for the mass distribution. At the same time, this leads to a less wide Gaussian peak and to the fact that small values of $M_{\text{max}}$ are now disfavoured. We also recover some minor changes in the posteriors for $d$, $\alpha_Z$ and $\gamma_Z$. We estimate that the expected number of events is $N_{\text{exp}} = 41_{-5}^{+6}$, which matches the 41 events used.

APPENDIX B: PARAMETER ESTIMATION INCLUDING GW190521

The posteriors of all parameters when including the event GW190521 can be seen in Fig. B1. Again all of the posteriors are in agreement with the ones obtained excluding GW190521. We find a slightly less steep power-law for the $m_1$ distribution with respect to the run excluding GW190521 with a power-law index for the mass distribution of $\alpha = 3.0_{-0.7}^{+0.9}$. We also retrieve a flatter $\sigma_g$ posterior and a posterior for $M_{\text{max}}$ that disfavors smaller values of masses, though fully in agreement with the previous results. These differences are because GW190521 components are very massive. For the model to fit the extra support at high masses, it needs a less steep power-law index for the mass distribution. At the same time, this leads to a less wide Gaussian peak and to the fact that small values of $M_{\text{max}}$ are now disfavoured. We also recover some minor changes in the posteriors for $d$, $\alpha_Z$ and $\gamma_Z$. We estimate that the expected number of events is $N_{\text{exp}} = 35_{-5}^{+6}$, which matches the 35 events used.

APPENDIX C: HIGH MASS MODEL

Here we employ a new model to be able to fit higher masses and we refer to it as the High mass model. Instead of imposing the window function $W_d$ to the PLG peak, we instead leave the power-law intact and only impose the windowing to the Gaussian peak of the distribution (see Fig. C1). The posteriors obtained by this model can be seen in Fig. C2. It is apparent that most of the posteriors are in agreement with those obtained by the usual model. However, the posterior for $M_{\text{max}}$ seems to give more stringent constraints in this case, with high values of masses being disfavoured. This is expected since in this model the power law is left intact allowing for significant support for higher values in the mass distribution after the $M_{\text{PSN}}$. Therefore, because we do not have posterior samples in these ranges of masses, the Bayesian inference can exclude the very high mass values. This does not affect our estimations for the time delay parameters, since the posteriors are in agreement with those from our main model.
Figure A1. Posterior distributions for all the hyperparameters. Here we have fixed all cosmological parameters at Planck-2018 cosmology. Different posteriors for different SNR thresholds can be seen.
**Figure B1.** Posterior distributions for all the hyperparameters. Here we have fixed all cosmological parameters at Planck-2018 cosmology. Different posteriors including or excluding GW190521 can be seen.
Figure C1. Comparison of the $m_1$ distribution for different mass models. For the values of parameters, we selected the median values of our estimations from Case-2 at $z_m = 0.1$.

Figure C2. Posterior distributions for all the hyperparameters. Here we have fixed all cosmological parameters at Planck-2018 cosmology. The posteriors for the high mass model can be seen, as well as the posteriors obtained for the usual model.