

BAYESIAN LEARNING WITH WASSERSTEIN BARYCENTERS

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ABSTRACT. We introduce and study a novel model-selection strategy for Bayesian learning, based on optimal transport, along with its associated predictive posterior law: the Wasserstein population barycenter of the posterior law over models. We first show how this estimator, termed Bayesian Wasserstein barycenter (BWB), arises naturally in a general, parameter-free Bayesian model-selection framework, when the considered Bayesian risk is the Wasserstein distance. Examples are given, illustrating how the BWB extends some classic parametric and non-parametric selection strategies. Furthermore, we also provide explicit conditions granting the existence and statistical consistency of the BWB, and discuss some of its general and specific properties, providing insights into its advantages compared to usual choices, such as the model average estimator. Finally, we illustrate how this estimator can be computed using the stochastic gradient descent (SGD) algorithm in Wasserstein space introduced in a companion paper [7], and provide a numerical example for experimental validation of the proposed method.

INTRODUCTION

Given a set \mathcal{M} of probability distributions on some data space X , learning a model $m \in \mathcal{M}$ from data points $D = (x_1, \dots, x_n) \in X^n$ consists in choosing, under a given criterion, an element of \mathcal{M} that *best* explains D as data sampled from it. The Bayesian paradigm provides a probabilistic approach to deal with model uncertainty in terms of a *prior distribution* on models \mathcal{M} , and also furnishes strategies to address the problem of model selection, based on the *posterior distribution* on \mathcal{M} given D . This type of estimators, usually called *predictive posterior laws*, include classical Bayesian estimators such as the maximum a posteriori estimator (MAP), the posterior mean, the Bayesian model average estimator (BMA) and generalizations thereof. Predictive posterior estimators typically result from selection criteria consisting in optimizing some loss function, averaged with respect to the posterior law over models, or Bayesian risk function. We refer the reader to [23, 37] and references therein for mathematical background on Bayesian statistics and their use in the machine learning community.

In Section 1, we will formulate the general problem of Bayesian model selection directly on the space of probability measures (or models) on the data space, and show how this abstract framework covers both classic finitely-parametrized settings and parameter-free model spaces, allowing us to retrieve classical selection criteria as particular cases. An eye-opening observation that will follow from adopting this viewpoint is that many classical predictive posteriors can be seen as instances of *Fréchet means* [22], or barycenters in the space of probability measures, with respect to specific metrics or divergences between them, that play the role of abstract loss functions defined on the model space.

Building upon this general framework, the main goals of this work are to introduce a novel Bayesian model-selection criterion by proposing a loss function on models coming from the theory of optimal transport, and to study some of the distinctive features of the predictive posterior law that results from it. More precisely, let us consider observations $D = (x_1, \dots, x_n)$ in a metric space (X, d) and a set of candidate models \mathcal{M} that generated these observations. Equipping the set \mathcal{M} with a prior distribution Π , and denoting the corresponding posterior distribution over \mathcal{M} by Π_n , we will define the *Bayesian Wasserstein*

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barycenter estimator (BWB) as a minimizer $\hat{m}_p^n \in \mathcal{M}$ of the “risk function”

$$\mathcal{M} \ni m \mapsto \int_{\mathcal{P}(\mathcal{X})} W_p(m, \bar{m})^p \Pi_n(d\bar{m}), \quad (0.1)$$

where $\mathcal{P}(\mathcal{X})$ is the set of probability measures on \mathcal{X} , $p \geq 1$ and W_p is the celebrated p -Wasserstein distance on probability measures on \mathcal{X} associated with d , see [45, 46].

Wasserstein barycenters were initially studied in [1] and, since then, the concept has been extensively explored from both theoretical and practical perspectives. We refer the reader to the overview [38] for statistical applications, to the works [15, 16, 41, 12] for applications in machine learning, and to [10, 33, 3, 4] for a presentation of recent developments and further references. As a cautionary tale, we mention that the problem of computing Wasserstein barycenters is known to be NP-hard, see [2]. To cope with this, fast methods aiming at learning and generating approximate Wasserstein barycenters on the basis of neural networks techniques, have also been proposed, see e.g. [32, 31].

In Section 2 we will recall Wasserstein distances and revisit Wasserstein barycenters together with their basic properties such as existence, uniqueness and absolute continuity. In Section 3, we will rigorously introduce the BWB estimator \hat{m}_p^n , which will correspond to the so-called *population Wasserstein barycenter* [33] for the posterior distribution on models Π_n , and we will state some of its main properties. Specifically, we will show that the BWB has less variance than the BMA and we will study its statistical consistency. In particular, we will address the question of “posterior consistency”, or asymptotic concentration of the posteriors Π_n around the Dirac mass on a model m_0 as $n \rightarrow \infty$, whenever the data consists of i.i.d. observations following the law m_0 , alongside the question of convergence of the BWB to the “true” distribution m_0 of the data in that setting. We refer the reader to [17, 23], and references therein, for detailed accounts on posterior consistency, a highly desirable feature of a Bayesian estimation procedure, both from a semi-frequentist perspective as well as from the “merging of opinions” point of view on Bayesian statistics (cf. [23, Chapter 6]). After reviewing central notions and tools from the framework of posterior consistency, namely the celebrated Schwartz’ theorem [44], [23, Theorem 6.17] and the notion of Kullback-Leibler support of Π , we will provide in that section equivalent and (verifiable) sufficient conditions for both the posterior consistency in the Wasserstein topology and the a.s. convergence

$$\lim_{n \rightarrow \infty} W_p(\hat{m}_p^n, m_0) = 0 \text{ a.s.}$$

to hold, when \hat{m}_p^n is the BWB computed with n i.i.d. observations sampled from m_0 .

Additionally, Sections 1 to 3 present a series of examples illustrating the main concepts of our work, their relationship to standard objects in Bayesian statistics and the applicability of our theoretical results.

Lastly, we will show how the BWB estimator can be calculated using a novel stochastic algorithm, introduced in the companion paper [7], to compute population Wasserstein barycenters in a general setting. This algorithm, presented in Section 4, can be seen as an abstract *stochastic gradient descent method in the Wasserstein space* and is advantageous compared to gradient or fixed-point algorithms developed in [3, 4, 48], whose application is restricted to barycenters of model spaces \mathcal{M} comprised of finitely-many elements only. Moreover, our algorithm has theoretical guarantees of convergence under suitable conditions, it can be easily implemented for some families of regular models for which optimal transport maps are explicit or easily computed (we recall one such family in Section 5.1), and its convergence rate can be studied and established in some cases, see [13] for the Gaussian setting. A comprehensive numerical experiment illustrating this method, and its natural “batch” variants, will be presented in Section 5 and compared to (more conventional) empirical barycenter estimators.

Notation:

- We denote $\mathcal{P}(\mathcal{X})$ the set of (Borel) probability measures on \mathcal{X} endowed with the weak topology, and $\mathcal{P}_{ac}(\mathcal{X})$ the (measurable) subset of absolutely continuous probability measures, with respect to a common reference σ -finite measure λ on \mathcal{X} .

- As a convention, we shall use the same notation for an element $m(dx) \in \mathcal{P}_{ac}(\mathcal{X})$ and its density $m(x)$ with respect to λ .
- We denote by $\text{supp}(\nu)$ the support of a measure ν and by $|\text{supp}(\nu)|$ its cardinality.
- Given $\Gamma \in \mathcal{P}(\mathcal{P}(\mathcal{X}))$ and a measurable subset $\mathcal{M} \subseteq \mathcal{P}(\mathcal{X})$, we say that \mathcal{M} is a *model space* for Γ if $\Gamma(\mathcal{M}) = 1$.
- Last, given a measurable map $T : \mathcal{Y} \rightarrow \mathcal{Z}$ and a measure ν on \mathcal{Y} we denote by $\nu_{\#}T$ the image measure (or push-forward), that is, the measure on \mathcal{Z} given by $\nu_{\#}T(\cdot) = \nu(T^{-1}(\cdot))$.

1. BAYESIAN LEARNING IN MODEL SPACE

We start by setting a general framework for Bayesian learning which covers both finitely-parametrized settings (including hierarchical models) and parameter-free models. We consider a probability measure $\Pi \in \mathcal{P}(\mathcal{P}(\mathcal{X}))$ understood as a *prior* distribution on the model space $\mathcal{M} \subseteq \mathcal{P}_{ac}(\mathcal{X})$. In particular we have

$$\Pi(\mathcal{M}) = \Pi(\mathcal{P}_{ac}(\mathcal{X})) = 1.$$

For each $n \in \mathbb{N} \setminus \{0\}$, Π canonically induces a law $\mathbf{\Pi}$ on $\mathcal{X}^n \times \mathcal{M}$, representing the joint law of a random model m chosen according to Π and a sample

$$D = (x_1, \dots, x_n) \subseteq \mathcal{X}^n$$

of i.i.d. observations drawn from it. That is,

$$\mathbf{\Pi}(dx_1, \dots, dx_n, dm) := m(dx_1) \cdots m(dx_n) \Pi(dm) = m(x_1) \cdots m(x_n) \lambda(dx_1) \cdots \lambda(dx_n) \Pi(dm).$$

Note that, in the above equation and throughout, $\Pi(dm)$ denotes integration over $m \in \mathcal{M}$ w.r.t. $\Pi \in \mathcal{P}(\mathcal{P}(\mathcal{X}))$, whereas integration over $x \in \mathcal{X}$ w.r.t. $m \in \mathcal{P}(\mathcal{X})$ is denoted $m(dx) = m(x)\lambda(dx)$. The law on \mathcal{X}^n of the data D , conditionally on a model m , is thus given by

$$\mathbf{\Pi}(dx_1, \dots, dx_n | m) := m(x_1) \cdots m(x_n) \lambda(dx_1) \cdots \lambda(dx_n), \quad (1.1)$$

with density $\mathbf{\Pi}(x_1, \dots, x_n | m) = m(x_1) \cdots m(x_n)$ with respect to $\lambda^{\otimes n}$, and the marginal density of D with respect to $\lambda^{\otimes n}$ is $\mathbf{\Pi}(x_1, \dots, x_n) := \int_{\mathcal{M}} \bar{m}(x_1) \cdots \bar{m}(x_n) \Pi(d\bar{m})$.

The *posterior distribution* $\mathbf{\Pi}(dm | x_1, \dots, x_n)$ given the data D is also an element of $\mathcal{P}(\mathcal{P}(\mathcal{X}))$, which we denote Π_n for simplicity and which, by virtue of the Bayes rule, is given in this setting by

$$\Pi_n(dm) := \frac{\mathbf{\Pi}(x_1, \dots, x_n | m) \Pi(dm)}{\mathbf{\Pi}(x_1, \dots, x_n)} = \frac{m(x_1) \cdots m(x_n) \Pi(dm)}{\int_{\mathcal{M}} \bar{m}(x_1) \cdots \bar{m}(x_n) \Pi(d\bar{m})}. \quad (1.2)$$

Notice that $(x_1, \dots, x_n) \mapsto \Pi_n = \mathbf{\Pi}(\cdot | x_1, \dots, x_n)$ defines $\lambda^{\otimes n}$ a.e. a measurable function from \mathcal{X}^n to $\mathcal{P}(\mathcal{P}(\mathcal{X}))$. The density $\Lambda_n(m)$ of $\Pi_n(dm)$ with respect the prior $\Pi(dm)$ is called the *likelihood* function. The fact that $\Pi_n \ll \Pi$ implies that a model space \mathcal{M} for Π is a model space for Π_n too.

We call *loss function* a non-negative functional on models $L : \mathcal{M} \times \mathcal{M} \rightarrow \mathbb{R}$, interpreting $L(m_0, \bar{m})$ as the cost of selecting model $\bar{m} \in \mathcal{M}$ when the true model is $m_0 \in \mathcal{M}$. With a loss function and the posterior distribution over models Π_n , the Bayes risk (or expected loss) $R_L(\bar{m} | D)$ and the corresponding Bayes estimator \hat{m}_L (or predictive posterior law) are respectively defined as follows:

$$R_L(\bar{m} | D) := \int_{\mathcal{M}} L(m, \bar{m}) \Pi_n(dm), \quad (1.3)$$

$$\hat{m}_L \in \operatorname{argmin}_{\bar{m} \in \mathcal{M}} R_L(\bar{m} | D). \quad (1.4)$$

See [8] for further background on Bayes risk and statistical decision theory. A key consequence of defining both L and Π_n directly on the model space \mathcal{M} (rather than on parameter space), is that learning according to eqs. (1.3)-(1.4) does not depend on the chosen parametrization or the geometry of the parameter space. Moreover, this point of view will allow us to define loss functions in terms of various metrics/divergences directly on the space $\mathcal{P}(\mathcal{X})$, and therefore to enhance the classical Bayesian estimation framework through the use of optimal transportation distances on that space. Before further developing these

ideas, we discuss how this general framework includes model spaces which are finitely parametrized, and recall some standard choices in that setting. We will also discuss the advantages of formulating the problem of model selection directly on the model space, even when this space can be finitely parametrized.

1.1. Parametric setting. We say that \mathcal{M} is finitely parametrized if there is an integer k , a measurable set $\Theta \subseteq \mathbb{R}^k$ termed parameter space, and a measurable function $\mathcal{T} : \Theta \mapsto \mathcal{P}_{ac}(\mathcal{X})$, called parametrization, such that $\mathcal{M} = \mathcal{T}(\Theta)$. In other words, $m = \mathcal{T}(\theta)$ is the model corresponding to parameter θ , which is classically denoted $p(\cdot|\theta)$ or $p_\theta(\cdot)$. In general, \mathcal{T} is a one-to-one function (the model space is otherwise said to be *over-parametrized*).

In the standard parametric Bayesian framework, a prior distribution is a probability law $p \in \mathcal{P}(\Theta)$ over Θ , typically assumed to have a (equally denoted) density $p(\theta)$ with respect to the Lebesgue measure. A function $\ell : \Theta \times \Theta \rightarrow \mathbb{R}_+$ is called a loss function (on parameters), whereby $\ell(\theta_0, \bar{\theta})$ is interpreted as the cost of choosing parameter $\bar{\theta}$ when the true parameter is θ_0 . The parametric Bayes risk [8] of $\bar{\theta} \in \Theta$ is then given by

$$R_\ell(\bar{\theta}|D) = \int_{\Theta} \ell(\theta, \bar{\theta}) p(\theta|x_1, \dots, x_n) d\theta, \quad (1.5)$$

where $p(\theta|x_1, \dots, x_n)$ is the posterior density of θ given observations x_1, \dots, x_n . The associated Bayes estimator is defined as $\hat{\theta}_\ell \in \operatorname{argmin}_{\bar{\theta} \in \Theta} R_\ell(\bar{\theta}|D)$. Hence, if the model space \mathcal{M} is finitely parametrized, learning a model boils down to finding the *best* model parameter $\theta \in \Theta$ under a given criterion, quantified by the parametric risk function $R_\ell(\cdot|D)$. Among continuous-valued losses on the parameter space, the *de facto* choice is the quadratic one $\ell_2(\theta, \bar{\theta}) = \|\theta - \bar{\theta}\|^2$, whose associated Bayes estimator is the posterior mean $\hat{\theta}_{\ell_2} = \int_{\Theta} \theta p(\theta|D) d\theta$. For one-dimensional parameter spaces, the absolute loss $\ell_1(\theta, \bar{\theta}) = |\theta - \bar{\theta}|$ yields the posterior median(s) estimator(s). The 0-1 loss formally given by $\ell_{0-1}(\theta, \bar{\theta}) = 1 - \delta_{\bar{\theta}}(\theta)$, with $\delta_{\bar{\theta}}$ the Dirac mass at $\bar{\theta}$, yields the risk $R_{\ell_{0-1}}(\bar{\theta}|D) = 1 - p(\bar{\theta}|D)$ and its corresponding Bayes estimator is the posterior mode or maximum a posteriori estimator (MAP), $\hat{\theta}_{\ell_{0-1}} = \hat{\theta}_{MAP}$.

The parametric case is embedded into the considered model-space setting as follows. The push-forward of p through \mathcal{T} defines a prior $\Pi = p_{\#}\mathcal{T}$ over the model space \mathcal{M} in the sense discussed at the beginning of this section. If \mathcal{T} is one to one, a loss function $\ell : \Theta \times \Theta \rightarrow \mathbb{R}_+$ induces a loss function L on the model space $\mathcal{M} = \mathcal{T}(\Theta)$, such that $L(\mathcal{T}(\theta_0), \mathcal{T}(\bar{\theta})) = \ell(\theta_0, \bar{\theta})$. More generally, any loss function L on $\mathcal{M} \times \mathcal{M}$ induces a loss functional ℓ on $\Theta \times \Theta$ defined as $\ell(\theta_0, \bar{\theta}) := L(\mathcal{T}(\theta_0), \mathcal{T}(\bar{\theta}))$. Moreover, when the data x_1, \dots, x_n under the model parameterized by θ consists of an i.i.d. sample from $p(\cdot|\theta) = \mathcal{T}(\theta)$, one can verify that Π_n given in eq. (1.2) corresponds precisely to the the push-forward through \mathcal{T} of $p(\theta|x_1, \dots, x_n)$ and that $R_\ell(\bar{\theta}|D)$ in eq. (1.5) is given by

$$R_\ell(\bar{\theta}|D) = \int_{\mathcal{M}} L(m, \bar{m}) \Pi_n(dm),$$

with $\Pi_n(dm) = \Lambda_n(m) \Pi(dm)$ associated with the prior on model space $\Pi = p_{\#}\mathcal{T}$, and $\bar{m} = \mathcal{T}(\bar{\theta})$.

Example 1.1. Consider the parametric Bayesian model with parameter space Θ and sample space \mathcal{X} both equal to \mathbb{R}^d and Gaussian parametrized models $\mathcal{N}(\theta, \Sigma)$, with $\Sigma \in \mathbb{R}^{d \times d}$ a fixed covariance matrix, and θ a random mean with prior $\mathcal{N}(\mu_0, \Sigma_0)$. The mean $\mu_0 \in \mathbb{R}^d$ and covariance matrix Σ_0 are fixed hyperparameters. This is classically denoted:

$$p(x|\theta) = \mathcal{N}(x; \theta, \Sigma), \quad p(\theta) = p(\theta|\mu_0, \Sigma_0) = \mathcal{N}(\theta; \mu_0, \Sigma_0).$$

From now on, $\mathcal{N}(y; \nu, \mathcal{K})$ stands for the density of the Gaussian law $\mathcal{N}(\nu, \mathcal{K})$ evaluated on the value y . Following from the introduced notation, the parametrization $\mathcal{T} : \Theta \rightarrow \mathcal{P}(\mathcal{X})$ is thus given by $\mathcal{T}(\theta) = \mathcal{N}(\theta, \Sigma)$, the model space is $\mathcal{M} = \{\mathcal{N}(\theta, \Sigma) : \theta \in \mathbb{R}^d\}$ and a measure m sampled from the prior $\Pi(dm) = p_{\#}\mathcal{T}(dm)$ is a Gaussian distribution on $\mathcal{X} = \mathbb{R}^d$, with fixed covariance matrix Σ and random mean θ distributed according to $\mathcal{N}(\mu_0, \Sigma_0)$. In this case, if both Σ and Σ_0 are nonsingular, the posterior of θ given $D = (x_1, \dots, x_n)$ is

$$p(\theta|D) = p(\theta|D, \mu_0, \Sigma_0) = \mathcal{N}\left(\theta; (\Sigma_0^{-1} + n\Sigma^{-1})^{-1}(\Sigma_0^{-1}\mu_0 + n\Sigma^{-1}\bar{x}_n), (\Sigma_0^{-1} + n\Sigma^{-1})^{-1}\right), \quad (1.6)$$

with \bar{x}_n the sample mean of the observations D . A model m sampled from the posterior $\Pi_n(dm)$ given D is thus obtained by sampling θ distributed according to $p(\theta|D)$ in eq. (1.6) and then setting $m = \mathcal{N}(\theta, \Sigma)$. The posterior mean estimator of the parameter θ is therefore given by

$$\hat{\theta}_{\ell_2} = (\Sigma_0^{-1} + n\Sigma^{-1})^{-1}(\Sigma_0^{-1}\mu_0 + n\Sigma^{-1}\bar{x}_n), \quad (1.7)$$

and the corresponding predictive posterior model is the Gaussian law $\mathcal{N}(\hat{\theta}_{\ell_2}, \Sigma)$. Observe that, in this case, one could equivalently obtain $\hat{\theta}_{\ell_2}$ as the mean of the predictive posterior law \hat{m}_L associated with the loss function L on models, defined by $L(m, \bar{m}) = \|\int_{\mathbb{R}^d} x m(dx) - \int_{\mathbb{R}^d} x \bar{m}(dx)\|^2$. This illustrates the equivalence of (some) pairs of losses in the model and parameter spaces that result on the same Bayes estimators. Additionally, observe that Bayesian inference on the mean and the covariance matrix could be similarly formulated in terms of a loss functions on models too, take e.g. $L'(m, \bar{m}) = L(m, \bar{m}) + \|\Sigma_m^{1/2} - \Sigma_{\bar{m}}^{1/2}\|_{\text{Fro}}^2$ with $L(m, \bar{m})$ as before, Σ_m the covariance matrix of a r.v. with law m and $\|\cdot\|_{\text{Fro}}$ the Frobenius norm on matrices.

The general model space framework applies equally to models with parameter and data spaces that might be of different nature:

Example 1.2. Assume $\Theta = \mathbb{R}_+$ and $\mathcal{X} = \mathbb{N}$, with $p(x|\lambda) = e^{-\lambda} \frac{\lambda^x}{x!}$, $x \in \mathcal{X}$, and $p(\lambda) \propto e^{-\beta\lambda} \lambda^{\alpha-1}$, $\lambda \in \Theta$. In this case, $\mathcal{T} : \Theta \rightarrow \mathcal{P}(\mathcal{X})$ is given by $\mathcal{T}(\lambda) = \text{Pois}(\lambda)$, the Poisson distribution of parameter λ , and a measure m sampled from the prior $\Pi(dm) = p_{\sharp} \mathcal{T}(dm)$ is a Poisson law with random parameter λ distributed according to the Gamma(α, β) law. The latter is a conjugate prior for the Poisson distribution, and m sampled from $\Pi_n(dm)$, the posterior on models given $D = (x_1, \dots, x_n)$, is again a Poisson law $\text{Pois}(\lambda)$, with random parameter λ distributed according to Gamma($n\bar{x}_n + \alpha, n + \beta$).

Remark 1.3. Hierarchical models are also catered for in the proposed setting. For instance, if in Example 1.1 the hyperparameter μ_0 is random with known density π_0 on \mathbb{R}^d , it can be integrated out and the prior p on parameters becomes an infinite Gaussian mixture:

$$p(\theta) = p(\theta|\Sigma_0) = \int \mathcal{N}(\theta; \mu_0, \Sigma) \pi_0(\mu_0) d\mu_0.$$

The parametrization mapping \mathcal{T} is in this case the same as before, and the corresponding prior $\Pi(dm) = p_{\sharp} \mathcal{T}(dm)$ and posteriors $\Pi_n(dm)$ on models $m \in \mathcal{P}(\mathcal{X})$ follow the same rationale as above.

As a cautionary note, the following example illustrates how defining estimators directly in terms of parameters might result in non-intrinsic criteria for model selection.

Example 1.4. On the parameter space $\Theta = [0, 1]$ consider the priors $p(d\theta) = d\theta$ and $\hat{p}(d\theta) = 2\theta d\theta$, and their associated parametrization maps $\mathcal{T}(\theta) := B(\theta)$ and $\hat{\mathcal{T}} := B(\theta^2)$ respectively. Here $B(\xi)$ denotes the law of a $\{0, 1\}$ -valued Bernoulli r.v. with ξ the probability of it being equal to 1. Notice that we have $\Pi := p_{\sharp} \mathcal{T} = \hat{p}_{\sharp} \hat{\mathcal{T}}$, as the law of θ^2 under \hat{p} is uniform on Θ . Starting from the prior p , the posterior density of θ given observations x_1, \dots, x_N is proportional to $\theta^{S_N} (1 - \theta)^{N - S_N}$, with $S_N = |\{i \leq N : x_i = 1\}|$. Thus the MAP estimator for θ is in this case $\theta_N := \frac{S_N}{N}$. On the other hand, starting from the prior \hat{p} , the posterior density of θ is proportional to $\theta^{2S_N+1} (1 - \theta^2)^{N - S_N}$ and now the associated MAP estimator for θ is $\hat{\theta}_N := \sqrt{\frac{2S_N+1}{2N+1}}$. Hence $\mathcal{T}(\theta_N) \neq \hat{\mathcal{T}}(\hat{\theta}_N)$ in general (although their discrepancy vanishes in the limit as $N \rightarrow \infty$). To summarize, although the same prior and posteriors at the level of models can arise by considering different parametrizations, the latter may easily define very different estimated models even if we agree on the estimation method (here the MAP).

1.2. Non parametric setting: posterior average estimators. The general ‘‘learning in model space’’ approach relies on loss functions that compare directly distributions (instead of their parameters), and thus allows us to define selection criteria based on intrinsic features of the models. It also allows for a wider choice of model-selection criteria, which can

account for geometric or information-theoretic properties of the models. The next result illustrates the fact that many examples of Bayesian estimators or predictive posterior, including the classical *model average estimator*, correspond to finding an instance of *Fréchet mean* or barycenter [22, 48] under a suitable metric/divergence on probability measures. See Appendix A for the proof.

Proposition 1.5. *Consider on the model $\mathcal{M} = \mathcal{P}_{ac}(\mathcal{X})$ the loss functions $L(m, \bar{m})$ given by:*

- i) *The L_2 -distance: $L(m, \bar{m}) = L_2(m, \bar{m}) := \frac{1}{2} \int_{\mathcal{X}} (m(x) - \bar{m}(x))^2 \lambda(dx)$,*
- ii) *The squared Hellinger distance $L(m, \bar{m}) = H^2(m, \bar{m}) := \frac{1}{2} \int_{\mathcal{X}} (\sqrt{m(x)} - \sqrt{\bar{m}(x)})^2 \lambda(dx)$.*
- iii) *The forward Kullback-Leibler divergence: $L(m, \bar{m}) = D_{KL}(m||\bar{m}) := \int_{\mathcal{X}} m(x) \ln \frac{m(x)}{\bar{m}(x)} \lambda(dx)$,*
- iv) *The reverse Kullback-Leibler divergence $L(m, \bar{m}) = D_{RKL}(m||\bar{m}) = D_{KL}(\bar{m}||m) := \int_{\mathcal{X}} \bar{m}(x) \ln \frac{\bar{m}(x)}{m(x)} \lambda(dx)$.*

Assume in each case that the infimum of the corresponding Bayes risk $\bar{m} \mapsto R_L(\bar{m}|D)$ defined in eq. (1.3) is attained with finite value in \mathcal{M} . Then, in cases i) and iii) the corresponding Bayes estimators (1.4) coincide with the standard Bayesian model average:

$$\hat{m}_{BMA}(x) := \mathbb{E}_{\Pi_n}[m](x) = \int_{\mathcal{M}} m(x) \Pi_n(dm). \quad (1.8)$$

Furthermore, the Bayes estimators corresponding to the cases ii) and iv) are given by the square model average and the exponential model average, respectively:

$$\hat{m}_{sqr}(x) = \frac{1}{Z_{sqr}} \left(\int_{\mathcal{M}} \sqrt{m(x)} \Pi_n(dm) \right)^2, \quad \hat{m}_{exp}(x) = \frac{1}{Z_{exp}} \exp \int_{\mathcal{M}} \ln m(x) \Pi_n(dm), \quad (1.9)$$

where Z_{sqr} and Z_{exp} denote the corresponding normalizing constants.

Example 1.6. If the posterior distribution was approximately equally concentrated on the models $m_0 = \mathcal{N}(\mu_0, 1)$ and $m_1 = \mathcal{N}(\mu_1, 1)$ with $\mu_0 \neq \mu_1$, that is, two (unimodal) Gaussian distributions with unit variance, then the standard model average is a bimodal non-Gaussian distribution with variance strictly larger than 1.

Example 1.7. In the parametric Bayesian model discussed in Example 1.1, the Bayesian model average estimator is the convolution of distributions on \mathbb{R}^d :

$$\begin{aligned} \hat{m}_{BMA}(x) &= \int_{\mathbb{R}^d} \mathcal{N}(x; \theta, \Sigma) \mathcal{N}\left(\theta; (\Sigma_0^{-1} + n\Sigma^{-1})^{-1}(\Sigma_0^{-1}\mu_0 + n\Sigma^{-1}\bar{x}_n), (\Sigma_0^{-1} + n\Sigma^{-1})^{-1}\right) d\theta \\ &= \int_{\mathbb{R}^d} \mathcal{N}(x - \theta; 0, \Sigma) \mathcal{N}\left(\theta; (\Sigma_0^{-1} + n\Sigma^{-1})^{-1}(\Sigma_0^{-1}\mu_0 + n\Sigma^{-1}\bar{x}_n), (\Sigma_0^{-1} + n\Sigma^{-1})^{-1}\right) d\theta \\ &= \mathcal{N}\left(x; (\Sigma_0^{-1} + n\Sigma^{-1})^{-1}(\Sigma_0^{-1}\mu_0 + n\Sigma^{-1}\bar{x}_n), \Sigma + (\Sigma_0^{-1} + n\Sigma^{-1})^{-1}\right), \end{aligned} \quad (1.10)$$

that is, a Gaussian density with mean equal to the posterior mean estimator $\hat{\theta}_{\ell_2}$ —see eq. (1.7)— and a covariance matrix that is strictly larger (in the usual order on nonnegative definite symmetric matrices) than that of the predictive posterior law $\mathcal{N}(\hat{\theta}_{\ell_2}, \Sigma)$ associated with it.

The Bayesian estimators considered Proposition 1.5, eqs. (1.8)-(1.9), share the following characteristic: their values at each point $x \in \mathcal{X}$ are computed in terms of some posterior *average* of the values of certain functions evaluated at x . This is due to the fact that the corresponding distances/divergences on probability distributions are “vertical” [43]: computing the distance between distributions m and \bar{m} involves the integration of vertical displacements between the graphs of their densities across their domain. An undesirable fact about *vertical averages* is that they are not well suited to incorporate *geometric* properties into the model space (as illustrated by Example 1.6). More generally, model averages might yield solutions that can be hardly interpretable in terms of the prior and parameters, or even be intractable. This motivates us to explore the use of “horizontal” distances between probability distributions, thus extending the concept of Bayes estimator by making use of the geometric features of the model space. The next section presents the main ideas we will rely on to build such Bayes estimators.

2. WASSERSTEIN DISTANCES AND BARYCENTERS: A QUICK REVIEW

We shall now introduce objects analogous to those in Proposition 1.5 but suited to *Wasserstein distances* on the space of probability measures. The following framework is adopted in the sequel:

Assumption 2.1. *The metric space (X, d) is a separable locally-compact geodesic space endowed with a σ -finite Borel measure λ and $p \geq 1$.*

By *geodesic* we mean that the space (X, d) is complete and any pair of points admit a mid-point with respect to d . Next, we briefly recall some basic elements of optimal transport theory and Wasserstein distances, referring to [45, 46] for general background.

2.1. Optimal transport and the Wasserstein distance. Given two measures μ, ν over X we denote by $\text{Cpl}(\mu, \nu)$ the set of transport plans or couplings with marginals μ and ν , i.e., $\gamma \in \text{Cpl}(\mu, \nu)$ if and only if $\gamma \in \mathcal{P}(X \times X)$, $\gamma(dx, X) = \mu(dx)$ and $\gamma(X, dy) = \nu(dy)$. Given a real number $p \geq 1$ we define the p -Wasserstein space $\mathcal{W}_p(X)$ by

$$\mathcal{W}_p(X) := \left\{ \eta \in \mathcal{P}(X) : \int_X d(x_0, x)^p \eta(dx) < \infty, \text{ some } x_0 \right\}.$$

The p -Wasserstein distance between measures μ and ν is given by

$$W_p(\mu, \nu) = \left(\inf_{\gamma \in \text{Cpl}(\mu, \nu)} \int_{X \times X} d(x, y)^p \gamma(dx, dy) \right)^{\frac{1}{p}}. \quad (2.1)$$

An optimizer of the right-hand side of eq. (2.1) always exists and is called an *optimal transport*. The distance W_p turns $\mathcal{W}_p(X)$ into a complete metric space. If in eq. (2.1) we assume that $p = 2$, X is the Euclidean space, and if μ is absolutely continuous, then Brenier's theorem [45, Theorem 2.12(ii)] establishes the uniqueness of a minimizer, and guarantees that it is supported on the graph of the subdifferential of a convex function. The corresponding gradient is thus called an *optimal transport map*. Explicit formulae for such optimal transport maps do exist in some cases, e.g., for generic one-dimensional distributions and multivariate Gaussians when $p = 2$ (see [14]). Contrary to the distances / divergences considered in Proposition 1.5, Wasserstein distances are *horizontal* [43], in the sense that they involve integrating horizontal displacements between the graphs of probability densities.

Example 2.2. The squared 2-Wasserstein distance between multivariate Gaussian distributions $m = \mathcal{N}(\theta, \Sigma)$, and $\bar{m} = \mathcal{N}(\bar{\theta}, \bar{\Sigma})$ is given by $W_2^2(m, \bar{m}) = \|\theta - \bar{\theta}\|^2 + \text{Tr} \left(\Sigma + \bar{\Sigma} - 2(\Sigma^{1/2} \bar{\Sigma} \Sigma^{1/2})^{1/2} \right)$ —see e.g. [24, 19]. Furthermore, if Σ and $\bar{\Sigma}$ commute, we have $W_2^2(m, \bar{m}) = \|\theta - \bar{\theta}\|^2 + \|\Sigma^{1/2} - \bar{\Sigma}^{1/2}\|_{\text{Fro}}^2$.

Example 2.3. The 1-Wasserstein distance between Poisson distributions $m = \text{Pois}(\lambda)$ and $\bar{m} = \text{Pois}(\bar{\lambda})$ is $|\lambda - \bar{\lambda}|$, which can be verified from the well-known expression $W_1(m, \bar{m}) = \int |m(-\infty, t] - \bar{m}(-\infty, t]| dt$ (valid for general one-dimensional distributions m, \bar{m}), and the so-called ‘‘Poisson-Gamma dual relation’’: $e^{-\lambda} \sum_{k=0}^n \frac{\lambda^k}{k!} = \int_{\lambda}^{\infty} \frac{t^n}{n!} e^{-t} dt$. When $\lambda > \bar{\lambda}$, the optimal coupling between $X \sim \text{Pois}(\lambda)$ and $Y \sim \text{Pois}(\bar{\lambda})$ is obtained taking $X \sim \text{Pois}(\lambda)$ and Y binomial with parameters $(X, \frac{\bar{\lambda}}{\lambda})$ conditionally on X . Alternatively, one could consider the coupling $(X, Y) := (N_\lambda, N_{\bar{\lambda}})$, where $\{N_t\}_{t \geq 0}$ is a Poisson process with intensity 1.

2.2. Wasserstein barycenter. Let us now recall the Wasserstein barycenter, introduced in [1] and further studied in [40, 28, 33], among others. Our definition slightly extends the ones in those works in that the optimization problem is posed in a possibly strict subset of the usual one.

Definition 2.4. *Let $\Gamma \in \mathcal{P}(\mathcal{P}(X))$. The p -Wasserstein risk of $\nu \in \mathcal{P}(X)$ is*

$$V_p(\nu) := \int_{\mathcal{P}(X)} W_p(m, \nu)^p \Gamma(dm) \leq +\infty.$$

Given a measurable set $\mathcal{M} \subseteq \mathcal{P}(\mathcal{X})$, any measure $\hat{m}_p \in \mathcal{M}$ which attains the quantity

$$\inf_{\nu \in \mathcal{M}} V_p(\nu)$$

with finite value, is called a p -Wasserstein barycenter of Γ over \mathcal{M} .

Notice that in principle we are not assuming \mathcal{M} to be a model space for Γ , but this will often be the case. When the support of Γ is infinite and $\mathcal{M} = \mathcal{W}_p(\mathcal{X})$, this object is termed p -Wasserstein population barycenter of Γ as introduced in [33]; see [10].

Example 2.5. Given two univariate Gaussian distributions $m_0 = \mathcal{N}(\mu_0, \sigma_0^2)$ and $m_1 = \mathcal{N}(\mu_1, \sigma_1^2)$, one can verify —using the expression in Example 2.2— that the 2-Wasserstein barycenter for $\Gamma(dm) = \frac{1}{2}\delta_{m_0}(dm) + \frac{1}{2}\delta_{m_1}(dm)$ is given by $\hat{m} = \mathcal{N}(\frac{\mu_0 + \mu_1}{2}, (\frac{\sigma_0 + \sigma_1}{2})^2)$. This should be compared to Example 1.6. Fig. 1 illustrates the corresponding vertical and a horizontal interpolations between two Gaussian densities with different means and the same variance.

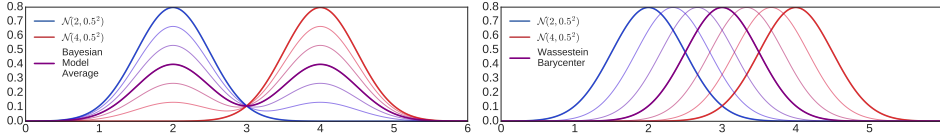


FIGURE 1. Vertical interpolation (left) and horizontal interpolation (right) of two Gaussian densities.

Let us introduce additional notation for the sequel and review then some basic properties of Wasserstein barycenters. Considering $\mathcal{W}_p(\mathcal{X})$ with the complete metric W_p as a base Polish metric space, we define $\mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$ in the natural way: $\Gamma \in \mathcal{P}(\mathcal{W}_p(\mathcal{X}))$ is an element of $\mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$ if it is concentrated on a set of measures with finite moments of order p , and moreover for some (and then all) $\tilde{m} \in \mathcal{W}_p(\mathcal{X})$ it satisfies

$$\int_{\mathcal{P}(\mathcal{X})} W_p(m, \tilde{m})^p \Gamma(dm) < \infty.$$

We endow $\mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$ with the corresponding p -Wasserstein distance, which we also denote W_p for simplicity. Also, if Γ is concentrated on measures with finite moments of order p which have densities with respect to λ , then we write $\Gamma \in \mathcal{P}(\mathcal{W}_{p,ac}(\mathcal{X}))$ and use the notation $\Gamma \in \mathcal{W}_p(\mathcal{W}_{p,ac}(\mathcal{X}))$ if, as before, $\int_{\mathcal{P}(\mathcal{X})} W_p(m, \tilde{m})^p \Gamma(dm) < \infty$ for some \tilde{m} .

Remark 2.6. If $\Gamma \in \mathcal{P}(\mathcal{P}(\mathcal{X}))$ has a p -Wasserstein barycenter \hat{m}_p over \mathcal{M} , then

$$\infty > \int_{\mathcal{P}(\mathcal{X})} W_p(m, \hat{m}_p)^p \Gamma(dm) = \int_{\mathcal{W}_p(\mathcal{X})} W_p(m, \hat{m}_p)^p \Gamma(dm),$$

hence $\Gamma \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$. Moreover, $\Gamma \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$ is equivalent to the corresponding model average $\tilde{m}(dx) := \mathbb{E}_\Gamma[m](dx)$ having a finite p -moment, since for any $y \in \mathcal{X}$,

$$\int_{\mathcal{W}_p(\mathcal{X})} W_p(\delta_y, m)^p \Gamma(dm) = \int_{\mathcal{W}_p(\mathcal{X})} \int_{\mathcal{X}} d(y, x)^p m(dx) \Gamma(dm) = \int_{\mathcal{X}} d(y, x)^p \int_{\mathcal{W}_p(\mathcal{X})} m(dx) \Gamma(dm).$$

We next state an existence result first established in [33, Theorem 2] for the case $\mathcal{M} = \mathcal{W}_p$. See Appendix B for a simpler, more direct, proof.

Theorem 2.7. *Suppose Assumption 2.1 holds, $\Gamma \in \mathcal{P}(\mathcal{P}(\mathcal{X}))$, and $\mathcal{M} \subseteq \mathcal{W}_p$ is a weakly closed set. There exists a p -Wasserstein barycenter of Γ over \mathcal{M} if and only if $\Gamma \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$.*

Regarding uniqueness, the following general result was proven in [33, Proposition 6] for the case $\mathcal{X} = \mathbb{R}^q$ with d the Euclidean distance and $p = 2$ (observe that, in that situation, the previous result applies):

Lemma 2.8. *Assume $\Gamma \in \mathcal{W}_2(\mathcal{W}_2(\mathbb{R}^q))$ and that there exists a set $A \subseteq \mathcal{W}_2(\mathbb{R}^q)$ of measures with*

$$\mu \in A, B \in \mathcal{B}(\mathbb{R}^q), \dim(B) \leq q - 1 \implies \mu(B) = 0,$$

and $\Gamma(A) > 0$. Then, Γ admits a unique 2-Wasserstein population barycenter over $\mathcal{W}_2(\mathbb{R}^q)$.

Remark 2.9. Observe that the model space $\mathcal{M} = \mathcal{W}_{p,ac}(\mathcal{X})$ is not weakly closed. Nevertheless, the existence and uniqueness of a population barycenter over that set can still be guaranteed when $p = 2$, $\mathcal{X} = \mathbb{R}^q$, d is the Euclidean distance, λ is the Lebesgue measure, and

$$\Gamma(\{m : \|\frac{dm}{d\lambda}\|_\infty < \infty\}) > 0. \quad (2.2)$$

This was proven in [28, Theorem 6.2] for compact finite-dimensional manifolds with lower-bounded Ricci curvature (equipped with the volume measure), but one can read-off the (non-compact but flat) Euclidean case $\mathcal{X} = \mathbb{R}^q$ from the proof therein, in order to establish the absolute continuity of a barycenter over $\mathcal{W}_2(\mathbb{R}^q)$, in the setting of Lemma 2.8. If $|\text{supp}(\Gamma)| < \infty$ then eq. (2.2) can be relaxed to the condition $\Gamma(\{m : m \ll \lambda\}) > 0$, as shown in [1] or [28, Theorem 5.1].

The following statement, corresponding to [7, Lemma 3.1], provides a useful description of barycenters which generalizes a result proven in [3] when $|\text{supp}(\Gamma)| < \infty$.

Lemma 2.10. *Assume $p = 2$, $\mathcal{X} = \mathbb{R}^q$, $d = \text{Euclidean distance}$, $\lambda = \text{Lebesgue measure}$. Let $\Gamma \in \mathcal{W}_2(\mathcal{W}_2(\mathcal{X}))$ and $\tilde{\Gamma} \in \mathcal{W}_2(\mathcal{W}_{2,ac}(\mathcal{X}))$. There exists a jointly measurable function $\mathcal{W}_{2,ac}(\mathbb{R}^q) \times \mathcal{W}_2(\mathbb{R}^q) \times \mathbb{R}^q \ni (\mu, m, x) \mapsto T_\mu^m(x)$ which is $\mu(dx)\Gamma(dm)\tilde{\Gamma}(d\mu)$ -a.s. equal to the unique optimal transport map from μ to m at x . Furthermore, letting $\hat{\mu}$ be a barycenter of Γ , we have*

$$x = \int T_{\hat{\mu}}^m(x)\Gamma(dm), \hat{\mu}(dx) - a.s.$$

3. BAYESIAN WASSERSTEIN BARYCENTER AND STATISTICAL PROPERTIES

Building on the Wasserstein distance as a loss function on models, we arrive to the following central object of the article:

Definition 3.1. *Let us consider a prior $\Pi \in \mathcal{P}(\mathcal{P}(\mathcal{X}))$ with model space $\mathcal{M} \subseteq \mathcal{W}_{p,ac}(\mathcal{X})$ and data $D = (x_1, \dots, x_n)$ which determines Π_n as in eq. (1.2). We define the p -Wasserstein Bayes risk of $\bar{m} \in \mathcal{W}_p(\mathcal{X})$ and a Bayes Wasserstein barycenter (BWB) estimator \hat{m}_p^n over \mathcal{M} respectively as follows:*

$$V_p^n(\bar{m}|D) := \int_{\mathcal{P}(\mathcal{X})} W_p(m, \bar{m})^p \Pi_n(dm), \text{ and} \quad (3.1)$$

$$\hat{m}_p^n \in \underset{\bar{m} \in \mathcal{M}}{\text{argmin}} V_p^n(\bar{m}|D), \quad (3.2)$$

if the corresponding minimum is finite.

Example 3.2. In the setting of Example 1.1, the 2-Wasserstein loss function on Gaussian models (see Example 2.2) induces the usual quadratic loss on the mean parameters: $W_2^2(m, \bar{m}) = \ell(\theta, \bar{\theta}) := \|\theta - \bar{\theta}\|^2$. Thus, in the notation of eq. (1.5), we have

$$V_2^n(\bar{m}|D) = R_\ell(\bar{\theta}|D).$$

This implies that, for any tuple of data point $D = (x_1, \dots, x_n)$ the BWB corresponds to the Gaussian distribution $\mathcal{N}(\hat{\theta}, \Sigma) \in \mathcal{M}$ with mean $\hat{\theta} = \hat{\theta}_{\ell_2} = (\Sigma_0^{-1} + n\Sigma^{-1})^{-1}(\Sigma_0^{-1}\mu_0 + n\Sigma^{-1}\bar{x}_n)$, that is, the posterior mean estimator of the mean parameter. Moreover, in this particular case the barycentric cost or optimal 2-Wasserstein Bayes risk equals the trace of the covariance of a random vector with law given in eq. (1.6), i.e., $\text{Tr}((\Sigma_0^{-1} + n\Sigma^{-1})^{-1})$. Notice that the covariance Σ of the BWB is strictly smaller than that of the corresponding BMA estimator in eq. (1.10) (in the usual order on symmetric positive semidefinite matrices). This is, in fact, a general property of the BWB as claimed in Proposition 3.9 below.

Example 3.3. Similarly, in the parametric setting of Example 1.2, the problem of finding BWB estimators can be written in terms of the parametric loss induced by the corresponding p -Wasserstein distance on Poisson distributions, computed in Example 2.3. We thus deduce for $p = 1$ that the BWB estimator \hat{m}_1^n corresponds, in that setting, to the law $\text{Pois}(\lambda_1)$ with λ_1 the median of the posterior distribution $\text{Gamma}(n\bar{x}_n + \alpha, n + \beta)$; for which an explicit expression is not available.

Example 3.4. Let us assume $\mathcal{X} = \mathbb{R}$ and that $\Pi(dm)$ is supported on continuous models m over \mathbb{R} . Let F_m and $Q_m := F_m^{-1}$ denote, respectively, the cumulative distribution function and the right-continuous quantile function of such m . The coupling $(x, T_{m_0}^m(x))$, with x distributed like m_0 and $T_{m_0}^m$ the increasing map $T_{m_0}^m(x) := Q_m(F_{m_0}(x))$, is known to be optimal for the p -Wasserstein distance, for any $p \geq 1$ (see [45, Remark 2.19(iv)]). The BWB $\hat{m}_n = \hat{m}_n^p$ of the posterior Π_n is also independent of p and is characterized via its quantile, as follows:

$$Q_{\hat{m}_n}(\cdot) = \int_{\mathcal{M}} Q_m(\cdot) \Pi_n(dm).$$

Interestingly enough, the model average $\bar{m}_n := \int m \Pi_n(dm)$ of Π_n is in turn characterized by its averaged cumulative distribution function: $F_{\bar{m}_n}(\cdot) = \int_{\mathcal{M}} F_m(\cdot) \Pi_n(dm)$. See Section 5 in the companion paper [7] for details and further discussion on one-dimensional Wasserstein barycenters, in particular, on geometric properties they inherit from the elements m of the support of the prior/posterior.

Remark 3.5. In the sequel, unless otherwise stated, an a.s. statement about Π_n is meant to hold almost surely with respect to the marginal law $m^{\otimes n}(dx_1, \dots, dx_n) \int_{\mathcal{M}} \Pi(dm)$ of a data sample of size n (sometimes called prior predictive distribution). In particular, for $\Pi(dm)$ —almost every m , such statement holds for $m^{\otimes n}(dx_1, \dots, dx_n)$ —almost every sample (x_1, \dots, x_n) .

Remark 3.6. We observe that $\Pi \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$ implies for each $n \geq 1$ that

$$\Pi_n \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X})) \text{ a.s.}$$

Indeed, for fixed $\tilde{m} \in \mathcal{W}_p(\mathcal{X})$, (we thank an anonymous referee for pointing this out) the following quantity is finite:

$$\begin{aligned} \int_{\mathcal{M}} W_p(m, \tilde{m})^p \Pi(dm) &= \int_{\mathcal{X}^n} \int_{\mathcal{M}} W_p(m, \tilde{m})^p m(x_1) \cdots m(x_n) \Pi(dm) \lambda(dx_1) \cdots \lambda(dx_n) \\ &= \int_{\mathcal{X}^n} \int_{\mathcal{M}} W_p(m, \tilde{m})^p \Pi_n(dm) \int_{\mathcal{M}} \tilde{m}(x_1) \cdots \tilde{m}(x_n) \Pi(d\tilde{m}) \lambda(dx_1) \cdots \lambda(dx_n) \\ &= \int_{\mathcal{X}^n} \left[\int_{\mathcal{M}} W_p(m, \tilde{m})^p \Pi_n(dm) \right] \tilde{m}^{\otimes n}(dx_1, \dots, dx_n) \int_{\mathcal{M}} \Pi(d\tilde{m}). \end{aligned} \tag{3.3}$$

However, $\Pi \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$ is in general not enough for $\Pi_n \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$ to hold a.s. for i.i.d. data points (x_1, \dots, x_n) sampled from a *fixed* law m_0 , which would be the natural setting to formulate the question of Bayesian consistency (see next subsection). Corollary 3.19 below ensures this fact for suitable given laws m_0 , in the framework of Bayesian consistency. In Appendix C we further provide a sufficient condition on the prior Π (termed *integrability after updates*) ensuring that $\Pi_n \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$ for *every* possible tuple of data points (x_1, \dots, x_n) .

The following statement gathers the discussion in Section 2.2 for the case $\Gamma = \Pi_n$, as well as the main point of Remark 3.6:

Theorem 3.7. *Suppose Assumption 2.1 holds and the model space \mathcal{M} is weakly closed. Let $\Pi \in \mathcal{P}(\mathcal{P}(\mathcal{X}))$ be a prior with model space $\mathcal{M} \subseteq \mathcal{W}_{p,ac}(\mathcal{X})$ and Π_n be the corresponding posterior given the data $D = (x_1, \dots, x_n)$. The following are equivalent:*

- a) A p -Wasserstein barycenter estimator \hat{m}_p^n for Π_n over \mathcal{M} exists a.s.

- b) $\Pi_n \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$ a.s.
- c) The model average $\bar{m}^n(dx) = \mathbb{E}_{\Pi_n}[m](dx)$ has a.s. a finite p -moment.

Moreover, if $\Pi \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$, a p -Wasserstein barycenter estimator \hat{m}_p^n over \mathcal{M} exists a.s. for every $n \geq 1$.

We thus make in all the sequel the following assumption:

Assumption 3.8. $\Pi \in \mathcal{W}_p(\mathcal{W}_{p,ac}(\mathcal{X}))$ and there exists a weakly closed model space $\mathcal{M} \subseteq \mathcal{W}_{p,ac}(\mathcal{X})$ for Π .

Since $\Pi_n \ll \Pi$ a.s., Assumption 3.8 together with Remark 3.6 imply $\Pi_n \in \mathcal{W}_p(\mathcal{W}_{p,ac}(\mathcal{X}))$ a.s. for all n .

We will next study some basic statistical properties of the BWB estimator.

3.1. Variance reduction with respect to BMA. In this subsection, we assume that $\mathcal{X} = \mathbb{R}^q$, $\lambda =$ Lebesgue measure, $d =$ Euclidean distance and $p = 2$. Let $\hat{m} := \hat{m}_2^n$ be the unique population barycenter of Π_n in that case, and denote by $(m, x) \mapsto T^m(x)$ a measurable function equal $\lambda(dx)\Pi(dm)$ a.e. to the unique optimal transport map from \hat{m} to $m \in \mathcal{W}_2(\mathcal{X})$. As a consequence of Lemma 2.10 we have the fixed-point property $\hat{m} = (\int T^m \Pi(dm))(\hat{m})$. Thus, for all convex functions φ , non negative or with at most quadratic growth, we have

$$\begin{aligned} \mathbb{E}_{\hat{m}}[\varphi(x)] &= \int_{\mathcal{X}} \varphi(x) \hat{m}(dx) = \int_{\mathcal{X}} \varphi \left(\int_{\mathcal{M}} T^m(x) \Pi_n(dm) \right) \hat{m}(dx) \\ &\leq \int_{\mathcal{X}} \int_{\mathcal{M}} \varphi(T^m(x)) \Pi_n(dm) \hat{m}(dx) = \int_{\mathcal{M}} \int_{\mathcal{X}} \varphi(T^m(x)) \hat{m}(dx) \Pi_n(dm) \\ &= \int_{\mathcal{M}} \int_{\mathcal{X}} \varphi(x) m(dx) \Pi_n(dm) = \int_{\mathcal{X}} \varphi(x) \int_{\mathcal{M}} m(dx) \Pi_n(dm) = \mathbb{E}_{\hat{m}_{BMA}}[\varphi(x)], \end{aligned}$$

where $\hat{m}_{BMA} = \mathbb{E}_{\Pi_n}[m]$ is the Bayesian model average in eq. (1.8). We have used Jensen's inequality and Fubini's theorem. This means that the BWB estimator is less spread, or *smaller*, in the sense of convex-order of probability measures, than the BMA. As a consequence, we have:

Proposition 3.9. Consider $p = 2$, and let \hat{m}_{BMA} and \hat{m}_2^n respectively denote the BMA and the BWB estimators associated with Π_n . Then, we have $\mathbb{E}_{\hat{m}_{BMA}}[x] = \mathbb{E}_{\hat{m}_2^n}[x]$ and $\mathbb{E}_{\hat{m}_{BMA}}[\|x\|^2] \geq \mathbb{E}_{\hat{m}_2^n}[\|x\|^2]$. In other words, the BWB estimator has less variance than the BMA. Furthermore, with \bar{x} denoting the mean w.r.t. the BMA or BWB, the corresponding covariances satisfy: $\mathbb{E}_{\hat{m}_{BMA}}[(x - \bar{x})(x - \bar{x})^t] \geq \mathbb{E}_{\hat{m}_2^n}[(x - \bar{x})(x - \bar{x})^t]$ in the usual order for symmetric positive semidefinite matrices. The inequalities are strict unless Π_n is a Dirac mass.

Proof. Given the previous discussion, the equality of means is obtained by taking $\varphi(x)$ and $-\varphi(x)$ equal to each coordinate of x and the inequality of variances by taking $\varphi(x) = \|x\|^2$. The inequality for the covariances follows by taking $\varphi(x) = (y^t(x - \bar{x}))^2$ for arbitrary $y \in \mathbb{R}^d$.

For the last claim, we just need to make explicit the corresponding equality case of Jensen's inequality as used in the preceding discussion. Since $\|x\|^2$ is strictly convex, the equality of second moments implies that $\hat{m}_2^n(dx)$ a.e. x , that $T^m(x)$ is $\Pi_m(dm)$ a.s. constant. This entails that the map $T := T^m$ does not depend on m and that for $\Pi_m(dm)$ a.e. m , it holds that $m = T(\hat{m}_2^n)$. The equality case for the covariances is reduced to the previous one considering their traces using also the equality of means. \square

3.2. Convergence to the true model and Bayesian consistency. A natural question in Bayesian statistics is whether a given predictive posterior estimator is *consistent* (see [44, 17, 23]). In short, this means the convergence of the predictive posterior law, in some specified sense, towards the *true* model m_0 , as we observe more and more i.i.d. data sampled from m_0 . We are specifically interested in the question of whether the BWB estimator converges to the model m_0 and, more precisely, on conditions which guarantee that

$$W_p(\hat{m}_p^n, m_0) \rightarrow 0, \quad m_0^{(\infty)} \text{ a.s.} \quad (3.4)$$

as $n \rightarrow \infty$, where \hat{m}_p^n is for each n a BWB over a model space \mathcal{M} . Here and in the sequel, $m_0^{(\infty)}$ denotes the product law of the infinite sample $\{x_n\}_n$ of i.i.d. data distributed according to m_0 . We will see that this question is linked to the notion of consistency (cf. [23, Definition 6.1]) of the prior, introduced next:

Definition 3.10. *The prior Π is said to be consistent at m_0 in the weak topology (resp. p -Wasserstein topology) if for each open neighbourhood U of m_0 in the weak topology of $\mathcal{P}(\mathcal{X})$ (resp. p -Wasserstein topology of $\mathcal{W}_p(\mathcal{X})$), one has $\Pi_n(U^c) \rightarrow 0$, $m_0^{(\infty)}$ - a.s.*

Remark 3.11. We notice that in the literature of Bayesian consistency, see e.g. [44, 23], Π satisfying the above property (in either topology) would be called “strongly consistent” at m_0 in allusion to the $m_0^{(\infty)}$ -a.s. convergences, whereas the term “weakly consistent” would be used when those convergences hold in probability. Since we will be only dealing with the almost sure notion, in order to avoid possible confusions with topological concepts, the adverb “strongly” is omitted throughout when referring to consistency of the prior, while the adverb “weakly” only refers to the weak topology on probability measures.

The celebrated Schwartz theorem [44] provides sufficient conditions for consistency w.r.t. a given topology, see also [23, Chapter 6] for a modern treatment. A key ingredient is the notion of Kullback-Leibler support:

Definition 3.12. *A measure m_0 is an element the Kullback-Leibler support of Π , denoted*

$$m_0 \in \text{KL}(\Pi),$$

if $\Pi(m : D_{\text{KL}}(m_0||m) < \varepsilon) > 0$ for every $\varepsilon > 0$, with $D_{\text{KL}}(m_0||m)$ the reverse Kullback-Leibler entropy defined as $\int \log \frac{m_0}{m}(x)m_0(dx)$ if $m_0 \ll m$ and as $+\infty$ otherwise.

Remark 3.13. The statistical model is interpreted as being “correct” or *well specified*, if the data distribution m_0 is an element of $\text{supp}(\Pi)$, the support of Π w.r.t. the weak topology, see [9, 26, 29, 30, 23]. The condition $m_0 \in \text{KL}(\Pi)$ is stronger. Indeed, by the Csiszar-Pinsker inequality and the fact that the dual bounded-Lipschitz distance (metrizing the weak topology in $\mathcal{P}(\mathcal{X})$) is majorized by the total variation distance, one can check that $\text{KL}(\Pi) \subseteq \text{supp}(\Pi)$. In particular, one has $\text{KL}(\Pi) \subseteq \mathcal{M}$ for any weakly closed model space \mathcal{M} for Π . (The reader may consult the mentioned works for the *misspecified* framework too.)

Remark 3.14. If $m_0 \in \text{KL}(\Pi)$, then $m_0(dx) \ll m(dx) \int_{\mathcal{M}} \Pi(dm)$, the marginal law of one data point x in the Bayesian model defined by Π . Indeed, for any measurable set $A \subseteq \mathcal{X}$ such that $\int_{\mathcal{M}} \int_A m(dx) \Pi(dm) = 0$ we have $\Pi(m : m_0 \ll m, m(A) = 0) = \Pi(m : m_0 \ll m) \geq \Pi(m : D_{\text{KL}}(m_0||m) < +\infty) > 0$. This will be useful later.

We recall (see Theorem 6.17 and Example 6.20 in [23]) the following result which concerns the weak topology:

Theorem 3.15. *Assume (only) that (\mathcal{X}, d) is Polish endowed with a σ -finite Borel measure λ , that $\Pi \in \mathcal{P}_{ac}(\mathcal{X})$ and that $m_0 \in \text{KL}(\Pi)$. Then, Π is consistent at m_0 in the weak topology.*

We now state a general result relating consistency of Π at m_0 in the p -Wasserstein topology and the convergence (3.4) (or consistency of the predictive posterior \hat{m}_p^n), with other asymptotic properties of the posterior laws in the Wasserstein setting. Recall that the notation W_p throughout stands for the Wasserstein distance both in $\mathcal{W}_p(\mathcal{X})$ and $\mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$.

Theorem 3.16. *Suppose Assumptions 2.1 and 3.8 hold, that $m_0 \in \mathcal{M}$, and that*

$$\Pi_n \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X})), \quad m_0^{\otimes n} - \text{ a.s. for all } n \geq 1. \quad (3.5)$$

The following are equivalent:

- a) $W_p(\Pi_n, \delta_{m_0}) \rightarrow 0$, $m_0^{(\infty)}$ -a.s. as $n \rightarrow \infty$.
- b) $m_0^{(\infty)}$ -a.s. as $n \rightarrow \infty$, we have $W_p(\hat{m}_p^n, m_0) \rightarrow 0$ and the barycentric cost (or optimal p -Wasserstein Bayes risk or) $\int_{\mathcal{M}} W_p(m, \hat{m}_p^n)^p \Pi_n(dm)$ goes to 0.

- c) Π is consistent at m_0 in the p -Wasserstein topology and the p -moment of the BMA estimator (1.8) converge $m_0^{(\infty)}$ -a.s. to that of m_0 as $n \rightarrow \infty$, i.e. for some (and then all) $x_0 \in \mathcal{X}$ we have

$$\int_{\mathcal{X}} d(x, x_0)^p \hat{m}_{BMA}^n(dx) = \int_{\mathcal{M}} \int_{\mathcal{X}} d(x, x_0)^p m(dx) \Pi_n(dm) \rightarrow \int_{\mathcal{X}} d(x, x_0)^p m_0(dx), m_0^{(\infty)} \text{ a.s.} \quad (3.6)$$

- d) Π is consistent at m_0 in the weak topology and for some (and then all) $x_0 \in \mathcal{X}$ we have

$$\int_{\mathcal{M}} \left| \int_{\mathcal{X}} d(x, x_0)^p m(dx) - \int_{\mathcal{X}} d(x, x_0)^p m_0(dx) \right| \Pi_n(dm) \rightarrow 0, m_0^{(\infty)} \text{ -a.s.} \quad (3.7)$$

Proof. By minimality of a barycenter over \mathcal{M} ,

$$\int_{\mathcal{M}} W_p(m, \hat{m}_p^n)^p \Pi_n(dm) \leq \int_{\mathcal{M}} W_p(m, m_0)^p \Pi_n(dm) = W_p(\Pi_n, \delta_{m_0})^p.$$

Thus, for some $c > 0$ depending only on p ,

$$W_p(m_0, \hat{m}_p^n)^p \leq c \int_{\mathcal{M}} W_p(m, \hat{m}_p^n)^p \Pi_n(dm) + c \int_{\mathcal{M}} W_p(m, m_0)^p \Pi_n(dm) \quad (3.8)$$

$$\leq 2c W_p(\Pi_n, \delta_{m_0})^p \quad (3.9)$$

proving that a) \Rightarrow b). The converse b) \Rightarrow a) follows from

$$\int_{\mathcal{M}} W_p(m, m_0)^p \Pi_n(dm) \leq c \int_{\mathcal{M}} W_p(m, \hat{m}_p^n)^p \Pi_n(dm) + c W_p(\hat{m}_p^n, m_0)^p.$$

Let us now show that a) \Rightarrow c). The convergence $W_p(\Pi_n, \delta_{m_0}) \rightarrow 0$ implies (by the Portmanteau theorem) that $\limsup_{n \rightarrow \infty} \Pi_n(F) \leq \delta_{m_0}(F)$ for all closed sets F of $\mathcal{W}_p(\mathcal{X})$. Taking $F = U^c$ with U a neighborhood of m_0 yields the consistency of Π at m_0 in the p -Wasserstein topology. Moreover it implies that for each $\tilde{m} \in \mathcal{W}_p(\mathcal{X})$,

$$\int_{\mathcal{M}} W_p^p(m, \tilde{m}) \Pi_n(dm) \rightarrow \int_{\mathcal{M}} W_p^p(m, \tilde{m}) \delta_{m_0}(dm) = W_p^p(m_0, \tilde{m}), m_0^{(\infty)} \text{ -a.s.}$$

as $n \rightarrow \infty$. Choosing $\tilde{m} = \delta_{x_0}$, we have $W_p^p(m, \tilde{m}) = \int_{\mathcal{X}} d(x, x_0)^p m(dx)$ for any m , from where (3.6) follows.

We next prove that c) \Rightarrow a). The space $\mathcal{W}_p(\mathcal{X})$ being Polish, there is countable basis \mathcal{U} of open neighborhoods of m_0 such that $m_0^{(\infty)}$ -a.s., $\Pi_n(U^c) \rightarrow 0$ for all $U \in \mathcal{U}$. Thus, if $G \subseteq \mathcal{W}_p(\mathcal{X})$ is any open set such that $m_0 \in G$, for some $U \in \mathcal{U}$ we have $U \subseteq G$ and therefore $\liminf_{n \rightarrow \infty} \Pi_n(G) \geq \liminf_{n \rightarrow \infty} \Pi_n(U) = 1 - \lim_{n \rightarrow \infty} \Pi_n(U^c) = 1 = \delta_{m_0}(G)$, $m_0^{(\infty)}$ -a.s. This implies, by the Portmanteau theorem, that the sequence $(\Pi_n)_n$ weakly converges to δ_{m_0} , $m_0^{(\infty)}$ -a.s., as probability measures on the metric space $\mathcal{W}_p(\mathcal{X})$. As in the previous (converse) implication, we obtain from (3.6) the convergence of some moments of order p of Π_n , to the corresponding moment of δ_{m_0} , $m_0^{(\infty)}$ -a.s. This plus the weak convergence just established imply that $W_p(\Pi_n, \delta_{m_0}) \rightarrow 0$ $m_0^{(\infty)}$ -a.s.

We have thus established that a), b) and c) are equivalent. Notice now that the function

$$m \mapsto \Phi(m) := \left| \int_{\mathcal{X}} \left(d(x, x_0)^p - \int_{\mathcal{X}} d(y, x_0)^p m_0(dy) \right) m(dx) \right|, \quad (3.10)$$

is continuous on $\mathcal{W}_p(\mathcal{X})$, since $x \mapsto d(x, x_0)^p - \int_{\mathcal{X}} d(y, x_0)^p m_0(dy)$ is continuous with polynomial growth of order p on \mathcal{X} . Moreover, $|\Phi(m)| \leq \int_{\mathcal{X}} d(y, x_0)^p m_0(dy) + W_p^p(m, \delta_{x_0})$, that is, Φ has polynomial growth of order at most p on $\mathcal{W}_p(\mathcal{X})$. Therefore, if a) or equivalently c) holds, we have $m_0^{(\infty)}$ -a.s. that

$$\int_{\mathcal{M}} \Phi(m) \Pi_n(dm) \rightarrow \int_{\mathcal{M}} \Phi(m) \delta_{m_0}(dm) = \Phi(m_0) = 0$$

which is tantamount to (3.7). Moreover if c) holds, consistency of Π at m_0 in the weak topology is obvious. This shows that c) \Rightarrow d).

Last, if d) holds, we deduce with Markov's inequality that, for each rational $\varepsilon > 0$,

$$\Pi_n \left\{ m \in \mathcal{M} : \left| \int_{\mathcal{X}} d(x, x_0)^p m(dx) - \int_{\mathcal{X}} d(x, x_0)^p m_0(dx) \right| \geq \varepsilon \right\} \rightarrow 0, m_0^{(\infty)} \text{ a.s.} \quad (3.11)$$

as $n \rightarrow \infty$. This, together with the consistency of Π at m_0 w.r.t. weak topology, is equivalent to having that consistency w.r.t. the p -Wasserstein topology. Moreover, we have

$$\begin{aligned} & \left| \int_{\mathcal{X}} d(x, x_0)^p \hat{m}_{BMA}^n(dx) - \int_{\mathcal{X}} d(x, x_0)^p m_0(dx) \right|_V \\ & \leq \int_{\mathcal{M}} \left| \int_{\mathcal{X}} d(x, x_0)^p m(dx) - \int_{\mathcal{X}} d(x, x_0)^p m_0(dx) \right| \Pi_n(dm) \end{aligned}$$

and so the convergence (3.7) implies the convergence (3.6). This and the previous show that d) implies c), concluding the proof. \square

Remark 3.17. The convergence (3.11) for each $\varepsilon > 0$ is exactly what one must add to consistency at m_0 of Π in the weak topology to obtain such consistency in the p -Wasserstein topology. However, the latter is only equivalent to the $m_0^{(\infty)}$ -a.s. weak convergence of Π_n to δ_{m_0} as measures on the metric space $\mathcal{W}_p(\mathcal{X})$ as $n \rightarrow \infty$, which is not enough to grant the $m_0^{(\infty)}$ -a.s. convergence in $\mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$ of Π_n to δ_{m_0} . Similarly, consistency at m_0 of Π in the p -Wasserstein topology cannot in general be obtained by adding only the convergence of moments 3.6 to the consistency at m_0 of Π in the weak topology. Of course, if the space \mathcal{X} is bounded, weak and p -Wasserstein topologies on it coincide, and consistency of Π at m_0 in the weak topology *implies* in that case the convergence (3.7) (since the function Φ in the previous proof is in that case continuous and bounded). Hence, in that case, all the equivalent properties in Theorem 3.16 are satisfied. The same is true if $\Pi(dm)$ -a.e. m is supported on a fixed bounded (weakly) closed set $\mathcal{Y} \subseteq \mathcal{X}$ (just replace \mathcal{X} by \mathcal{Y}).

An immediate consequence of the proof of Theorem 3.16 (cf. the estimate (3.8)) is the following bound, which can be used to obtain quantitative estimates for the average rate of convergence of the BWB, if the posterior and the Wasserstein distance between models are explicit enough:

Corollary 3.18. *Under the assumptions of Theorem 3.16, for some constant $c_p > 0$ we have*

$$\mathbb{E} \left(W_p(m_0, \hat{m}_p^n)^p \right) \leq c_p \mathbb{E} \left(W_p(\Pi_n, \delta_{m_0})^p \right).$$

The following result gathers sufficient conditions for consistency at m_0 of Π in the p -Wasserstein topology and convergence of the BWB.

Corollary 3.19. *Suppose Assumptions 2.1 and 3.8 hold, and moreover that $m_0 \in \text{KL}(\Pi)$. Then Π is consistent at m_0 in the weak topology and condition (3.5) holds. If moreover either condition (i) or (ii) below hold, then Π is consistent at m_0 in the p -Wasserstein topology and $W_p(\hat{m}_p^n, m_0) \rightarrow 0, m_0^{(\infty)}$ -a.s. as $n \rightarrow \infty$:*

- (i) *the convergence (3.7) holds;*
- (ii) *for some $q > p, m_0^{(\infty)}$ -a.s. the q moments of the BMA estimator are bounded uniformly in n .*

Proof. In view of Theorem 3.15, to prove the first claim we just need to prove that (3.5) holds. To that end, let us first check that, whenever $m_0 \in \text{KL}(\Pi)$, we have $m_0^{\otimes n}(dx_1, \dots, dx_n) \ll m^{\otimes n}(dx_1, \dots, dx_n) \int_{\mathcal{M}} \Pi(dm)$. If $A \subseteq \mathcal{X}^n$ measurable is such that $\int_A m^{\otimes n}(dx_1, \dots, dx_n) \int_{\mathcal{M}} \Pi(dm) = 0$, then for $\Pi(dm)$ -a.e. m and every $x_i \in \mathcal{X}, i = 2, \dots, n$, one has $m(\{x \in \mathcal{X} : (x, x_2, \dots, x_n) \in A\}) = 0$, from which we get $m_0(\{x \in \mathcal{X} : (x, x_2, \dots, x_n) \in A\}) = 0$ by Remark 3.14, hence $m_0^{\otimes n}(A) = 0$. We conclude noting that the set $A := \{(x_1, \dots, x_n) \in \mathcal{X}^n : \Pi_n \notin \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))\}$ has null $m^{\otimes n}(dx_1, \dots, dx_n) \int_{\mathcal{M}} \Pi(dm)$ -measure, by Remark 3.6.

Given the previous and part d) of Theorem 3.16, the convergence (3.7) immediately implies that Π is consistent at m_0 in the p -Wasserstein topology and that $W_p(\hat{m}_p^n, m_0) \rightarrow 0$ $m_0^{(\infty)}$ -a.s. as $n \rightarrow \infty$. To conclude the proof it is enough to show that the $m_0^{(\infty)}$ -a.s. uniform boundedness of the q moments of the BMA estimator for some $q > p$, implies, under the given assumptions, that said convergence (3.7) holds. Consider to that end the function Φ defined by (3.10) in the proof of Theorem 3.16. For each $\varepsilon > 0$ we have

$$\begin{aligned} \int_{\mathcal{M}} \Phi(m) \Pi_n(dm) &\leq \varepsilon + \int_{\mathcal{M} \cap \{\Phi(m) \geq \varepsilon\}} \Phi(m) \Pi_n(dm) \\ &\leq \varepsilon + \left(\int_{\mathcal{M}} \Phi(m)^{q/p} \Pi_n(dm) \right)^{p/q} (\Pi_n(m : \Phi(m) \geq \varepsilon))^{1-p/q} \end{aligned}$$

The assumption on the q -moments of the BWA estimators imply that $\sup_n \int_{\mathcal{M}} \Phi(m)^{q/p} \Pi_n(dm)$ is finite, $m_0^{(\infty)}$ -a.s., following from applying Jensen's inequality to the convex function $s \mapsto |s|^{q/p}$ on \mathbb{R} and the integral w.r.t. $m(dx)$ defining $\Phi(m)$. Since $\varepsilon > 0$ is arbitrary, it just remains to ensure that $\Pi_n(m : \Phi(m) \geq \varepsilon) \rightarrow 0$ $m_0^{(\infty)}$ -a.s. as $n \rightarrow \infty$ (that is, the convergence (3.11) holds). The elementary relations $t = t \wedge R + (t - R)_+$ and $(t - R)_+ \leq t \mathbf{1}_{\{t > R\}}$ for all $t, R \geq 0$ yield the bound

$$\begin{aligned} \Pi_n(m : \Phi(m) \geq \varepsilon) &\leq \Pi_n \left(m : \left| \int_{\mathcal{X}} R \wedge d(x, x_0)^p m(dx) - \int_{\mathcal{X}} R \wedge d(x, x_0)^p m_0(dx) \right| \geq \varepsilon/2 \right) \\ &\quad + \Pi_n \left(m : \int_{\{x \in \mathcal{X} : d(x, x_0)^p > R\}} d(x, x_0)^p [m(dx) + m_0(dx)] \geq \varepsilon/2 \right), \end{aligned}$$

where the first term on the r.h.s. goes $m_0^{(\infty)}$ -a.s. to 0 as $n \rightarrow \infty$, since $x \mapsto R \wedge d(x, x_0)^p$ is a bounded continuous function and Π is consistent at m_0 in the weak topology. The second term on the r.h.s. is bounded by

$$\frac{2}{\varepsilon} \left[\sup_{n \in \mathbb{N}} \int_{\mathcal{M}} \int_{\mathcal{X}} \mathbf{1}_{\{d(x, x_0)^p > R\}} d(x, x_0)^p m(dx) \Pi_n(dm) + \int_{\mathcal{X}} \mathbf{1}_{\{d(x, x_0)^p > R\}} d(x, x_0)^p m_0(dx) \right].$$

For each $\varepsilon > 0$, this expression can be made arbitrarily small by taking R large enough, since $m_0 \in \mathcal{W}_p(\mathcal{X})$ and because the uniform boundedness of the q -moments of the BWA estimators implies their p -moments are uniformly integrable. We conclude that $\Pi_n(m : \Phi(m) \geq \varepsilon) \rightarrow 0$ $m_0^{(\infty)}$ -a.s. as $n \rightarrow \infty$ as desired. \square

Example 3.20. Given Gaussian distributions $m_0 = \mathcal{N}(\theta_0, \Sigma)$ and $m = \mathcal{N}(\theta, \Sigma)$ on $\mathcal{X} = \mathbb{R}^d$, one can check that $D_{KL}(m_0 \| m) = \frac{1}{2}(\theta - \theta_0)' \Sigma^{-1}(\theta - \theta_0)$ and hence that $\text{KL}(\Pi) = \text{supp}(\Pi)$ in the parametric Bayesian model dealt with in Examples 1.1 and 3.2. Therefore, by Theorem 3.15, Π in those examples is consistent w.r.t. the weak topology at $m_0 = \mathcal{N}(\theta_0, \Sigma)$ for any $\theta_0 \in \mathbb{R}^d$. Moreover, the prior Π is easily seen to be in $\mathcal{W}_2(\mathcal{W}_2(\mathbb{R}^d))$, hence Lemma 3.19 ensures that $\Pi_n \in \mathcal{W}_2(\mathcal{W}_2(\mathbb{R}^d))$, $m_0^{\otimes n}$ -a.s. for all $n \geq 1$. This fact alternatively follows from the existence of a 2-Wasserstein barycenter for Π_n for any data points D , verified in this case. To verify consistency of Π in the 2-Wasserstein topology at any m_0 as well as convergence of the BWB, let us compute $W_2^2(\Pi_n, \delta_{m_0})$ and apply directly Theorem 3.16. Denoting by θ a r.v. with law $\mathcal{N}(\hat{\theta}_{\ell_2}, (\Sigma_0^{-1} + n\Sigma^{-1})^{-1})$, we find

$$\begin{aligned} W_2^2(\Pi_n, \delta_{m_0}) &= \int_{\mathcal{M}} W_2^2(m, m_0) \Pi_n(dm) \\ &= \mathbb{E}_{\theta}(\|\theta - \theta_0\|^2) \\ &= \|\theta_0 - \hat{\theta}_{\ell_2}\|^2 + \mathbb{E}_{\theta}(\|\hat{\theta}_{\ell_2} - \theta\|^2) \\ &= W_2^2(\hat{m}_2^n, m_0) + \int_{\mathcal{M}} W_2^2(m, \hat{m}_2^n) \Pi_n(dm), \end{aligned}$$

with $\mathbb{E}_\theta(\|\hat{\theta}_{\ell_2} - \theta\|^2) = \int_{\mathcal{M}} W_2^2(m, \hat{m}_2^n) \Pi_n(dm) = \text{tr}((\Sigma_0^{-1} + n\Sigma^{-1})^{-1}) \rightarrow 0$ as $n \rightarrow \infty$ and

$$\|\theta_0 - \hat{\theta}_{\ell_2}\|^2 = W_2^2(\hat{m}_2^n, m_0) = \|(\Sigma_0^{-1} + n\Sigma^{-1})^{-1}(\Sigma_0^{-1}\mu_0 + n\Sigma^{-1}\bar{x}_n) - \theta_0\|^2$$

which $m_0^{(\infty)}$ a.s. goes to 0 as $n \rightarrow \infty$ by the Law of Large Numbers, whenever the data $(x_n)_{n \geq 1}$ consists of an i.i.d. sample from the true model m_0 . Notice that the identity $W_2^2(\Pi_n, \delta_{m_0}) = W_2^2(\hat{m}_2^n, m_0) + \int_{\mathcal{M}} W_2^2(m, \hat{m}_2^n) \Pi_n(dm)$ found in this case, can be seen as a *bias-variance type decomposition* for the posterior law Π_n on $\mathcal{W}_2(\mathcal{X})$. In this case, one can readily see that $\mathbb{E}(W_2^2(\hat{m}_2^n, m_0)) \leq C/n$.

Example 3.21. Although the BWB estimator \hat{m}_1^n in the Poisson parametric Bayesian model of Example 1.2 is not explicit (see Example 3.3), we can still apply Theorem 3.16 a) to prove that it converges to the true model generating the data. Indeed, if $m_0 = \text{Pois}(\lambda_0)$, using the expression for the 1-Wasserstein distance between Poisson laws computed in Example 2.3 we see that

$$W_1(\Pi_n, \delta_{m_0}) = \int W_1(m, m_0) \Pi_n(dm) = \mathbb{E}|\lambda - \lambda_0| \leq \mathbb{E}(|\lambda - \lambda_0|^2)^{1/2},$$

with λ a r.v. with law $\text{Gamma}(n\bar{x}_n + \alpha, n + \beta)$. An elementary computation using the Gamma distribution's mean and variance shows, in this case, that

$$\mathbb{E}(|\lambda - \lambda_0|^2) = \frac{n\bar{x}_n + \alpha}{(n + \beta)^2} + \left[\frac{n\bar{x}_n + \alpha}{n + \beta} - \lambda_0 \right]^2,$$

which goes to zero $m_0^{(\infty)}$ a.s. as $n \rightarrow \infty$, whenever the data $(x_n)_{n \geq 1}$ is an i.i.d. sample from the true model m_0 . We deduce with Corollary 3.19 that $\mathbb{E}(W_1(m_0, \hat{m}_1^n)) \leq C/\sqrt{n}$.

As noticed earlier, consistency of Π at m_0 w.r.t. the p -Wasserstein topology is not enough to grant that $W_p(\Pi_n, \delta_{m_0}) \rightarrow 0$ $m_0^{(\infty)}$ -a.s.. But we will see next that this is true under a boundedness condition on the support of Π . Recall that $\text{supp}(\Pi)$ is said to be bounded in $\mathcal{W}_p(\mathcal{X})$ if

$$\text{diam}(\Pi) := \sup_{m, \bar{m} \in \text{supp}(\Pi)} W_p(m, \bar{m}) < \infty.$$

A typical example is the finitely parametrized case with compact parameter space Θ and continuous parametrization mapping $\mathcal{T} : \Theta \rightarrow \mathcal{W}_p(\mathcal{X})$. More generally, $\text{diam}(\Pi) < \infty$ amounts to Π being supported on a set of models with centered p -moments bounded by a constant. In particular, \mathcal{X} and the support of every $m \in \text{supp}(\Pi)$ may be unbounded, and still $\text{supp}(\Pi)$ be bounded. We have

Lemma 3.22. *If Π is consistent at m_0 in the p -Wasserstein topology and $\text{supp}(\Pi)$ is bounded in $\mathcal{W}_p(\mathcal{X})$, then $W_p(\Pi_n, \delta_{m_0}) \rightarrow 0$ $m_0^{(\infty)}$ -a.s. as $n \rightarrow \infty$.*

Proof. Let $\varepsilon > 0$ and $B = \{m : W_p(m, m_0) < \varepsilon\}$, then

$$\begin{aligned} W_p(\Pi_n, \delta_{m_0})^p &= \int_{\mathcal{M}} W_p(m, m_0)^p \Pi_n(dm) \\ &= \int_B W_p(m, m_0)^p \Pi_n(dm) + \int_{B^c} W_p(m, m_0)^p \Pi_n(dm) \\ &\leq \varepsilon^p + \int_{B^c} W_p(m, m_0)^p \Pi_n(dm). \end{aligned}$$

Since $\varepsilon > 0$ is arbitrary, we only need to check that the second term in the last line goes $m_0^{(\infty)}$ -a.s. to zero as $n \rightarrow \infty$. By consistency we have $\Pi_n(B^c) \rightarrow 0$, $m_0^{(\infty)}$ -a.s. as $n \rightarrow \infty$, and since $\text{supp}(\Pi_n) \subseteq \text{supp}(\Pi)$, we conclude that

$$\int_{B^c} W_p(m, m_0)^p \Pi_n(dm) \leq \text{diam}(\Pi)^p \Pi_n(B^c) \rightarrow 0, \quad m_0^{(\infty)}\text{-a.s.}$$

□

Remark 3.23. Π consistent at m_0 in the weak topology and $\text{supp}(\Pi)$ bounded in $\mathcal{W}_p(\mathcal{X})$ is in general not enough to obtain the conclusion of Lemma 3.22. To see this, write

$$W_p(\Pi_n, \delta_{m_0})^p = \int_{B_w} W_p(m, m_0)^p \Pi_n(dm) + \int_{B_w^c} W_p(m, m_0)^p \Pi_n(dm), \quad (3.12)$$

where B_w is a fixed small weak neighborhood of m_0 (e.g. one can use the bounded Lipschitz distance to build B_w). As in the proof of Lemma 3.22, the second term in the r.h.s. of (3.12) goes $m_0^{(\infty)}$ -a.s. to zero as $n \rightarrow \infty$ in this case too. However there is no reason why the term $\int_{B_w} W_p(m, m_0)^p \Pi_n(dm)$ should be small. The reason is that for the functional $m \mapsto \Psi(m) := W_p(m, m_0)^p$, even if bounded on $\mathcal{M} := \text{supp}(\Pi)$, there is no reason why $\Psi|_{B_w \cap \mathcal{M}}$ should be small (no matter how small B_w may be). Indeed, the statement “ $\Psi|_{B_w \cap \mathcal{M}}$ is small if B_w is small” would mean mathematically that the weak and the p -Wasserstein topologies coincide on \mathcal{M} locally around m_0 , but this is not true in general, even if \mathcal{M} is bounded¹. This should not be confused with the fact that, if \mathcal{M} equipped with the p -Wasserstein topology and metric is bounded, then on $\mathcal{W}_p(\mathcal{M})$ weak and p -Wasserstein convergence coincide. Consistency at m_0 of Π in the weak topology, on the other hand, is equivalent to $\Pi_n \rightarrow \delta_{m_0}, m_0^{(\infty)}$ -a.s. in the weak topology of $\mathcal{W}_p(\mathcal{M})$, when \mathcal{M} is equipped with the weak topology. This does not imply $\Pi_n \rightarrow \delta_{m_0}, m_0^{(\infty)}$ -a.s. in the weak topology of $\mathcal{W}_p(\mathcal{M})$, when \mathcal{M} is equipped with the p -Wasserstein topology, and hence provides another point of view as to why the l.h.s. of (3.12) does not go to zero without stronger assumptions.

The next result based on Schwartz theorem provides a (rather strong) condition ensuring that the equivalent properties in Theorem 3.16 hold. Unfortunately, we have not been able to prove such a result under more general assumptions.

Theorem 3.24. *Under Assumptions 2.1 and 3.8, suppose moreover that $m_0 \in \text{KL}(\Pi)$ and that, for some $\lambda_0 > 0$ and $x_0 \in \mathcal{X}$, one has*

$$\sup_{m \in \text{supp}(\Pi)} \int_{\mathcal{X}} e^{\lambda_0 d^p(x, x_0)} m(dx) < +\infty.$$

Then, Π is consistent at m_0 in the p -Wasserstein topology. Moreover, we have $W_p(\Pi_n, \delta_{m_0}) \rightarrow 0, m_0^{(\infty)}$ -a.s. and the BWB estimator is consistent in the sense that

$$W_p(\hat{m}_p^n, m_0) \rightarrow 0, m_0^{(\infty)} - a.s.$$

Before proving Theorem 3.24 some remarks on its assumptions and proof are in order.

Remark 3.25. The uniform control assumed on p -exponential moments implies, by Jensen’s inequality, that $\sup_{m \in \text{supp}(\Pi)} W_p^p(m, \delta_{x_0}) = \sup_{m \in \text{supp}(\Pi)} \int_{\mathcal{X}} d^p(x, x_0) m(dx) < +\infty$. By triangle inequality, this in turn implies that $\text{supp}(\Pi)$ is bounded.

Remark 3.26. The general picture of Bayesian consistency (including the misspecified case) parallels in several aspects Sanov’s large deviations theorem (see e.g. the Bayesian Sanov Theorem in [26, Theorem in 2.1] and references therein), and it similarly relies on exponential controls of (posterior) integrals. The proof of Theorem 3.24 follows the argument of [23, Example 6.20], where Theorem 3.15 above is obtained by combining Schwartz’ theorem ([23, Theorem 6.17]) with Hoeffding’s concentration inequality, to get uniform exponential controls of the posterior mass of complements of weak neighborhoods of m_0 , which are defined in terms of bounded random variables. A p -exponential moment control is what is needed to derive such concentration inequalities for unbounded random variables (e.g. moments), defining neighborhoods in the Wasserstein topology. Notice that finite p -exponential moments are also required for Sanov’s theorem to hold in the p -Wasserstein topology [47]. The uniform bound on exponential moments appears however too strong an assumption (it does not hold e.g. in the setting of Example 3.20). The question of relaxing that condition is left for future work.

Remark 3.27. In the *misspecified* framework dealt with in [9, 26, 29, 30, 23], and paralleling results applicable to the weak topology in those works, we expect the convergence $\hat{m}_p^n \rightarrow \text{argmin}_{m \in \mathcal{M}} D_{KL}(m_0 || m)$ w.r.t. W_p to hold, under suitable assumptions.

¹For instance, for $\mathcal{X} = \mathbb{R}$ and $p = 1$ we have that $\mathcal{M} := \{m_n := \frac{n-1}{n} \delta_0 + \frac{1}{n} \delta_n\}_{n \in \mathbb{N}} \cup \{\delta_0\}$ is 1-Wasserstein bounded, and yet $m_n \rightarrow \delta_0$ weakly but not in 1-Wasserstein topology.

Proof of Theorem 3.24. First we show that if U is any $\mathcal{W}_p(X)$ -neighbourhood of m_0 then $m_0^{(\infty)}$ -a.s. we have $\liminf_n \Pi_n(U) \geq 1$. According to Schwartz Theorem in the extended form [23, Theorem 6.17], under the assumption that $m_0 \in KL(\Pi)$, it is enough to find for each such U a sequence of measurable functions or “tests” $\varphi_n : \mathcal{X}^n \rightarrow [0, 1]$ such that

- (1) $\varphi_n(x_1, \dots, x_n) \rightarrow 0$, $m_0^{(\infty)}$ - a.s. and
- (2) $\limsup_n \frac{1}{n} \log \left(\int_{U^c} m^{\otimes n} (1 - \varphi_n) \Pi(dm) \right) < 0$.

First we will construct tests $\{\varphi_n\}_n$ that satisfy Point 1 and Point 2 above, over an appropriate subbase of neighbourhoods U , to finally extend those properties to general neighborhoods.

Recall that $\mu_k \rightarrow \mu$ in $W_p(X)$ iff for all continuous functions ψ on X with $|\psi(x)| \leq K(1 + d^p(x, x_0))$ for some $K \in \mathbb{R}_+$ it holds that $\int_X \psi(x) \mu_n(dx) \rightarrow \int_X \psi(x) \mu(dx)$; see [46]. Given such ψ and $\varepsilon > 0$ we define the open sets

$$U_{\psi, \varepsilon} := \left\{ m : \int_X \psi(x) m(dx) < \int_X \psi(x) m_0(dx) + \varepsilon \right\},$$

which form a sub-base for the p -Wasserstein neighborhood system at the distribution m_0 . We can assume that $K = 1$ by otherwise considering $U_{\psi/K, \varepsilon/K}$ instead. Given a neighborhood $U := U_{\psi, \varepsilon}$ of m_0 as above, we define the test functions

$$\varphi_n(x_1, \dots, x_n) = \begin{cases} 1 & \text{if } \frac{1}{n} \sum_{i=1}^n \psi(x_i) > \int_X \psi(x) m_0(dx) + \frac{\varepsilon}{2}, \\ 0 & \text{otherwise.} \end{cases}$$

By the law of large numbers, $m_0^{(\infty)}$ - a.s. : $\varphi_n(x_1, \dots, x_n) \rightarrow 0$, so Point 1 is verified. Point 2 is trivial if $r := \Pi(U^c) = 0$, so we assume from now on that $r > 0$. Thanks to the exponential moments control assumed on $\text{supp}(\Pi)$, for $\Pi(dm)$ a.e. m the random variable $Z = 1 + d^p(X, x_0)$ with $X \sim m$ has a moment-generating function $\mathcal{L}_m(t)$ which is finite for all $t \in [0, \lambda_0]$, namely

$$\mathcal{L}_m(t) := \mathbb{E} \left[e^{tZ} \right] = e^t \int_X e^{td^p(x, x_0)} m(dx) < +\infty.$$

We thus have the bounds

$$\int_X |\psi(x)|^k m(dx) \leq \mathbb{E} \left[Z^k \right] \leq k! \mathcal{L}_m(t) t^{-k}, \quad \forall \lambda_0 \geq t > 0,$$

for all $k \in \mathbb{N}$. Therefore, we may apply Bernstein’s inequality in the form of [36, Corollary 2.10] to the random variables $\{-\psi(x_i)\}_i$ under the measure $m^{(\infty)}$ on $\mathcal{X}^{\mathbb{N}}$, obtaining for any $\alpha < 0$ that

$$m^{(\infty)} \left(\sum_{i=1}^n \left[\psi(x_i) - \int_X \psi(x) m(dx) \right] \leq \alpha \right) \leq e^{-\frac{\alpha^2}{2(v-c\alpha)}},$$

with $v := 2n \mathcal{L}_m(t) t^{-2}$, $c := t^{-1}$, and $0 < t \leq \lambda_0$. Going back to the tests φ_n and using the definition of U^c we deduce that

$$\begin{aligned} \int_{U^c} m^{\otimes n} (1 - \varphi_n) \Pi(dm) &= \int_{U^c} m^{\otimes n} \left(\frac{1}{n} \sum_{i=1}^n \psi(x_i) \leq \int_X \psi(x) m_0(dx) + \frac{\varepsilon}{2} \right) \Pi(dm) \\ &\leq \int_{U^c} m^{\otimes n} \left(\frac{1}{n} \sum_{i=1}^n \psi(x_i) \leq \int_X \psi(x) m(dx) - \frac{\varepsilon}{2} \right) \Pi(dm) \\ &= \int_{U^c} m^{\otimes n} \left(\sum_{i=1}^n \left[\psi(x_i) - \int_X \psi(x) m(dx) \right] \leq -\frac{n\varepsilon}{2} \right) \Pi(dm) \\ &\leq \int_{U^c} \exp \left\{ -\frac{n\varepsilon^2}{2} \frac{t^2}{8 \mathcal{L}_m(t) + t\varepsilon} \right\} \Pi(dm) \\ &\leq r \exp \left\{ -\frac{n\varepsilon^2}{2} \frac{t^2}{8 \sup_{m \in \text{supp}(\Pi)} \mathcal{L}_m(t) + t\varepsilon} \right\}. \end{aligned}$$

Thanks to the uniform control of exponential moments on the support, we conclude as desired that

$$\limsup_n \frac{1}{n} \log \left(\int_{U^c} m^{\otimes n} (1 - \varphi_n) \Pi(dm) \right) \leq -\frac{t^2 \varepsilon^2}{16 \sup_{m \in \text{supp}(\Pi)} \mathcal{L}_m(t) + 2t\varepsilon} < 0.$$

Now, a general neighborhood U contains a finite intersection of, say, $N \in \mathbb{N}$ elements of the sub-base, i.e. $\bigcap_{i=1}^N U_{\psi_i, \varepsilon_i} \subseteq U$, so

$$\int_{U^c} m^{\otimes n} (1 - \varphi_n) \Pi(dm) \leq \sum_{i=1}^N \int_{U_{\psi_i, \varepsilon_i}^c} m^{\otimes n} (1 - \varphi_n) \Pi(dm).$$

Therefore we can conclude as in the sub-base case that Point 2 is verified. All in all, we have established that Π is p -Wasserstein consistent at m_0 . Thanks to Lemma 3.22 and the boundedness of $\text{supp}(\Pi)$ (see Remark 3.25) the last two claims are immediate. \square

4. BWB CALCULATION THROUGH DESCENT ALGORITHMS IN WASSERSTEIN SPACE

In this section we review some of the available methods to compute Wasserstein barycenters, and then explain how they can be used to calculate the BWB estimator. We will first survey the method proposed by Álvarez-Esteban, Barrio, Cuesta-Albertos and Matrán in [3, Theorem 3.6] and by Zemel and Panaretos in [48, Theorem 3, Corollary 2], applicable in the setting where the probability distribution Γ on the model space has finite support. This method is interpreted in [48] as a *gradient descent in the Wasserstein space*. We will then discuss the main aspects of the *stochastic gradient descent in Wasserstein space* (SGDW), introduced in our companion paper [7], building also upon the gradient descent idea. The latter method allows moreover the computation of population barycenters for a distribution Γ on the model space, using a streaming of random probability distributions sampled from it. In particular, we will recall the conditions established in [7] which ensure its convergence. The reader is referred to [41, 16, 34, 18] for alternative approaches to computing Wasserstein barycenters.

The two discussed methods can be easily implemented when explicit analytical expressions for optimal transport maps between distributions in the model space are available (see [7] or Section 5 for some examples). Moreover, in that case these methods can easily be coupled with sampling procedures (MCMC or others) for the posterior distribution Π_n in order to compute the BWB. We will see that the computation of the BWB through the SGDW has several advantageous features. In particular, when expressions for optimal transport maps are known, it can be done at nearly the same cost as the posterior sampling.

From now on we specialize Assumption 2.1, and make the following set of assumptions:

Assumption 4.1. $p = 2$, $\mathcal{X} = \mathbb{R}^q$, $d = \text{Euclidean metric}$, $\lambda = \text{Lebesgue measure}$. Furthermore, $\Gamma \in \mathcal{W}_2(\mathcal{W}_{p,ac}(\mathbb{R}^q))$ and there is a model space $\mathcal{M} \subseteq \mathcal{W}_{p,ac}(\mathbb{R}^q)$ for Γ which is weakly closed.

The following concept will be central in the “gradient-type” algorithms we consider:

Definition 4.2. We say that $\mu \in \mathcal{W}_{2,ac}(\mathbb{R}^q)$ is a Karcher mean of $\Gamma \in \mathcal{W}_2(\mathcal{W}_{2,ac}(\mathbb{R}^q))$ if

$$\mu\left(\left\{x : x = \int_{\mathcal{W}_2(\mathbb{R}^q)} T_\mu^m(x) \Gamma(dm)\right\}\right) = 1.$$

Remark 4.3. It is known that any 2-Wasserstein barycenter is a Karcher mean (c.f. [48]). However, the class of Karcher means is in general a strictly larger one, see [3]. For conditions ensuring uniqueness of Karcher means, see [48, 38] for the case when the support of Γ is finite and [10] for the case of an infinite support. In one dimension, the uniqueness of Karcher means holds without further assumptions. See [7] for a deeper discussion.

4.1. Gradient descent on Wasserstein space. Consider $\Gamma \in \mathcal{W}_2(\mathcal{W}_{p,ac}(\mathbb{R}^q))$ finitely supported: for some $m_i \in \mathcal{W}_{p,ac}(\mathbb{R}^q)$, $i = 1, \dots, L \in \mathbb{N}$, we have

$$\Gamma = \sum_{i \leq L} \lambda_i \delta_{m_i}.$$

Following [3] and [48], we define an operator over $\mathcal{W}_{p,ac}(\mathcal{X})$ by

$$G(m) := \left(\sum_{i=1}^L \lambda_i T_m^{m_i}\right)(m). \quad (4.1)$$

Starting from $\mu_0 \in \mathcal{W}_{p,ac}(\mathcal{X})$ one can then define the sequence

$$\mu_{n+1} := G(\mu_n), \text{ for } n \geq 0. \quad (4.2)$$

The next result proven in [3, Theorem 3.6] and independently in [48, Theorem 3, Corollary 2], establishes the convergence of the above sequence to a fixed-point of G , which is nothing other than a Karcher mean for Γ :

Proposition 4.4. *The sequence $\{\mu_n\}_{n \geq 0}$ in eq. (4.2) is tight and every weakly convergent subsequence of $\{\mu_n\}_{n \geq 0}$ converges in W_2 to an absolutely continuous measure in $\mathcal{W}_2(\mathbb{R}^q)$ which is a Karcher mean of Γ . If some m_i has a bounded density, and if there exists a unique Karcher mean \hat{m} , then \hat{m} is the Wasserstein barycenter of Γ and $W_2(\mu_n, \hat{m}) \rightarrow 0$.*

Panaretos and Zemel [48, Theorem 1] discovered that the sequence (4.2) can indeed be interpreted as a gradient descent (GD) scheme with respect to the *Riemannian-like* structure of the Wasserstein space $\mathcal{W}_2(\mathbb{R}^q)$. In fact, the functional on $\mathcal{W}_2(\mathbb{R}^q)$ given by

$$F(m) := \frac{1}{2} \sum_{i=1}^L \lambda_i W_2^2(m_i, m)$$

has a *Frechet derivative* at each point $m \in \mathcal{W}_{2,ac}(\mathbb{R}^q)$, given by

$$F'(m) = - \sum_{i=1}^L \lambda_i (T_m^{m_i} - I) = I - \sum_{i=1}^L \lambda_i T_m^{m_i} \in L^2(m),$$

where I is the identity map in \mathbb{R}^q . This means that for each such m , one has

$$\frac{F(\hat{m}) - F(m) - \int_{\mathbb{R}^q} \langle F'(m)(x), T_{\hat{m}}^m(x) - x \rangle m(dx)}{W_2(\hat{m}, m)} \rightarrow 0, \quad (4.3)$$

when $W_2(\hat{m}, m)$ goes to zero, by virtue of [5, Corollary 10.2.7]. It follows from Brenier's theorem [45, Theorem 2.12(ii)] that \hat{m} is a fixed point of G defined in eq. (4.1) if and only if $F'(\hat{m}) = 0$. The gradient descent sequence in Wasserstein space (GDW) with step γ starting from $\mu_0 \in \mathcal{W}_{2,ac}(\mathbb{R}^q)$ is defined by (c.f. [48])

$$\mu_{n+1} := G_\gamma(\mu_n), \text{ for } n \geq 0, \text{ where}$$

$$G_\gamma(m) := [I + \gamma F'(m)](m) = \left[(1 - \gamma)I + \gamma \sum_{i=1}^L \lambda_i T_m^{m_i} \right](m) = \left[I + \gamma \sum_{i=1}^L \lambda_i (T_m^{m_i} - I) \right](m),$$

and it coincides with the sequence in eq. (4.2) if $\gamma = 1$. These ideas serve as inspiration for the stochastic gradient descent iteration in the next part.

4.2. Stochastic gradient descent for population barycenters. We recall next the stochastic gradient descent sequence introduced in the companion paper [7], where the reader is referred to for details. Additionally to the assumptions in the previous part, we will also make use of an extra one introduced in [7]:

Assumption 4.5. *Γ has a W_2 -compact model space $K_\Gamma \subseteq \mathcal{W}_{2,ac}(\mathbb{R}^q)$. Moreover this set is geodesically convex: for every $\mu, \nu \in K_\Gamma$ and $t \in [0, 1]$, $((1-t)I + tT_\mu^\nu)(\mu) \in K_\Gamma$, with I the identity operator.*

In particular, under these assumptions, for each $\nu \in \mathcal{W}_2(\mathbb{R}^q)$ and $\Gamma(dm)$ a.e. m , there is a unique optimal transport map T_m^ν from m to ν and, by [33, Proposition 6], the 2-Wasserstein population barycenter is unique. We notice that, although strong at first sight, assumption 4.5 can be guaranteed in suitable parametric situations (e.g., Gaussian, or even the location scatter setting recalled in Section 5.1), or under moment and density constraints on the measures in K_Γ (e.g., under uniform bounds on their moments of order $2 + \varepsilon$ and their Boltzmann entropy, which are geodesically convex functionals, see [5]).

Definition 4.6. *Let $\mu_0 \in K_\Gamma$, $m_k \stackrel{iid}{\sim} \Gamma$, and $\gamma_k > 0$ for $k \geq 0$. We define the stochastic gradient descent in Wasserstein space sequence (SGDW) by*

$$\mu_{k+1} := \left[(1 - \gamma_k)I + \gamma_k T_{\mu_k}^{m_k} \right](\mu_k), \text{ for } k \geq 0. \quad (4.4)$$

The sequence is a.s. well-defined, as one can show by induction that $\mu_k \in \mathcal{W}_{2,ac}(\mathbb{R}^q)$ a.s. thanks to Assumption 4.5. The rationale for definition 4.4 is similar to that of Section 4.1, though now we wish to emphasize the population case: If we call now

$$F(\mu) := \frac{1}{2} \int_{\mathcal{W}_2(\mathbb{R}^q)} W_2^2(\mu, m) \Gamma(dm) \quad (4.5)$$

the functional minimized by a 2-Wasserstein barycenter, then we (formally at least) expect

$$F'(\mu)(x) = - \int_{\mathcal{W}_2(\mathbb{R}^q)} (T_\mu^m - I) \Gamma(dm)(x). \quad (4.6)$$

Hence, $(I - T_\mu^m)$ with $m \sim \Gamma$, is an unbiased estimator of $F'(\mu)$. This immediately suggests the stochastic descent sequence (4.4) introduced in [7], drawing inspiration from the classic SGD ideas [42].

Clearly μ is a Karcher mean for Γ iff $\|F'(\mu)\|_{L^2(\mu)} = 0$. Just like for the GD sequence, the SGD sequence is typically expected to converge to stationary points, or Karcher means in the present setting, rather than to minimisers. Next theorem provides sufficient conditions for the SGD sequence to a.s. converge to a Wasserstein barycenter, and is the main result of [7]. The following assumption on the steps γ_k , standard in the framework of SGD methods, is needed:

$$\sum_{k=1}^{\infty} \gamma_k^2 < \infty \quad \text{and} \quad \sum_{k=1}^{\infty} \gamma_k = \infty. \quad (4.7)$$

Theorem 4.7. *Suppose Assumptions 4.1 and 4.5, as well as conditions (4.7) hold. Furthermore, suppose that Γ admits a unique Karcher mean. Then, the SGD sequence $\{\mu_k\}_k$ in eq. (4.4) is a.s. convergent to the unique 2-Wasserstein barycenter $\hat{\mu}$ of Γ . Moreover, we have $\hat{\mu} \in K_\Gamma$.*

4.3. Batch stochastic gradient descent on Wasserstein space. We briefly recall how the variance of the SGD sequence can be reduced by using batches:

Definition 4.8. *Let $\mu_0 \in K_\Gamma$, $m_k^i \stackrel{iid}{\sim} \Gamma$, and $\gamma_k > 0$ for $k \geq 0$ and $i = 1, \dots, S_k$. The batch stochastic gradient descent (BSGD) sequence is given by*

$$\mu_{k+1} := \left[(1 - \gamma_k)I + \gamma_k \frac{1}{S_k} \sum_{i=1}^{S_k} T_{\mu_k}^{m_k^i} \right] (\mu_k). \quad (4.8)$$

The following two results, extracted from [7], justify the above definition: The first result states that this sequence is still converging, while the second one states that batches help *reducing noise*:

Proposition 4.9. *Under the assumptions of Theorem 4.7 the BSGD sequence $\{\mu_t\}_{t \geq 0}$ in eq. (4.8) converges a.s. to the 2-Wasserstein barycenter of Γ .*

Proposition 4.10. *The batch estimator for $F'(\mu)$ of batch size S , given by $-\frac{1}{S} \sum_{i=1}^S (T_\mu^{m_i} - I)$, has integrated variance*

$$\mathbb{V}[-\frac{1}{S} \sum_{i=1}^S (T_\mu^{m_i} - I)] := \int \text{Var}_{(m_i) \sim \Pi^{\otimes S}} \left[\frac{1}{S} \sum_{i=1}^S (T_\mu^{m_i}(x) - x) \right] \mu(dx) = O(\frac{1}{S}),$$

i.e. $\mathbb{V}[-\frac{1}{S} \sum_{i=1}^S (T_\mu^{m_i} - I)]$ decreasing linearly in the batch size.

4.4. Computation of the BWB. It is immediate to deduce a simple methodology based on the SGD algorithm, to compute the BWB estimator for general (finitely or infinitely supported) posterior laws $\Pi_n \in \mathcal{W}_2(\mathcal{W}_{2,ac}(\mathbb{R}^q))$.

We make the practical assumption that we are capable of generating independent models m_i from the posteriors Π_n (in the parametric setting, this can be done through efficient Markov Chain Monte Carlo (MCMC) techniques [6, 25, 11] or transport sampling procedures [20, 39, 27, 35]). On the theoretical side, we assume conditions 4.1 and 4.5 are satisfied by $\Gamma = \Pi$, the prior law on models, implying the posterior Π_n a.s. satisfies those conditions for all n too.

The proposed method can be sketched as follows:

- (1) Given a prior on models Π and data x_1, \dots, x_n , sample $\mu_0^{(n)} \sim \Pi_n$ and set $k = 0$.
- (2) Sample $m_k^{(n)}$ independent from $\mu_0^{(n)}, m_0^{(n)}, \dots, m_{k-1}^{(n)}$
- (3) Set

$$\mu_{k+1}^{(n)} := \left[(1 - \gamma_k)I + \gamma_k T_{\mu_k^{(n)}}^{m_k^{(n)}} \right] (\mu_k^{(n)}), \text{ for } k \geq 0.$$

- (4) Increase k by 1 and go to (2).

The algorithm can be run until $k \in \mathbb{N}$ large enough such that the squared 2–Wasserstein distance

$$W_2^2(\mu_{k+1}^{(n)}, \mu_k^{(n)}) = \gamma_k^2 \int_{\mathbb{R}^q} |x - T_{\mu_k^{(n)}}^{m_k^{(n)}}(x)|^2 d\mu_k^{(n)}(x)$$

between $\mu_{k+1}^{(n)}$ and $\mu_k^{(n)}$ is repeatedly smaller than some given positive threshold.

Moreover, under the assumptions of Theorem 3.24, Π is consistent in the W_2 -topology at the law m_0 of the data x_1, \dots, x_0 and then, for some (random) large enough $n, k \in \mathbb{N}$ and given $\varepsilon > 0$ one has $W_2(\mu_k^{(n)}, m_0) \leq W_2(\mu_k^{(n)}, \hat{m}_2^n) + W_2(\hat{m}_2^n, m_0) \leq \varepsilon$, where \hat{m}_2^n is the *BWB*. Notice that, besides the sequential generation of a finite i.i.d. sequence $\mu_0^{(n)}, m_0^{(n)}, \dots, m_k^{(n)} \sim \Pi_n$, at each step $k+1$ one only needs to compose a new transport map $\left[(1 - \gamma_k)I + \gamma_k T_{\mu_k^{(n)}}^{m_k^{(n)}} \right]$ with the transport map pushing forward $\mu_0^{(n)}$ to $\mu_k^{(n)}$, cumulatively constructed in the previous iterations. If an expression for each map $T_{\mu_k^{(n)}}^{m_k^{(n)}}$ is available, this can easily be done (and stored), specifying (only) the values of the lastly computed map on a pre-fixed grid.

The batch version SGWD can be implemented in a similar way to compute the *BWB* estimator.

Remark 4.11. An alternative, natural approach would be to sample, for given n a fixed number k of realizations $m_i \stackrel{\text{iid}}{\sim} \Pi_n$, $i = 1, \dots, k$, and compute, using the GDW algorithm, the Wasserstein barycenter $\hat{m}_2^{(n,k)}$ of the (finitely supported) empirical measure

$$\Pi_n^{(k)} := \frac{1}{k} \sum_{i=1}^k \delta_{m_i} \in \mathcal{W}_2(\mathcal{W}_{2,ac}(\mathbb{R}^q),$$

or 2-Wasserstein empirical barycenter of Π_n (see [10]). Indeed, by Varadarajan's theorem, conditionally on Π_n , a.s. $\Pi_n^{(k)}$ converges weakly as $k \rightarrow \infty$ to Π_n , and in the W_2 metric as soon as $\Pi_n \in \mathcal{W}_2(\mathcal{W}_{2,ac}(\mathbb{R}^q))$. Since under our assumptions Π_n a.s. has a unique 2-Wasserstein barycenter \hat{m}_2^n and, by [33, Theorem 3], $\hat{m}_2^{(n,k)}$ converges with respect to W_2 a.s. as $k \rightarrow \infty$ to it, we also get through this approach that

$$W_2(\hat{m}_2^{(n,k)}, m_0) \leq W_2(\hat{m}_2^{(n,k)}, \hat{m}_2^n) + W_2(\hat{m}_2^n, m_0) \leq \varepsilon$$

for some (random) large enough $n, k \in \mathbb{N}$. The clear disadvantage of this method is computational: besides generating k samples of Π_n , we need to additionally run a possibly large number of GDW steps to approximate $\hat{m}_2^{(n,k)}$, and we need to evaluate at each step k new transport maps (instead of one, for the SGDW). Moreover, if additional k' new samples from Π_n become available, we need to run the whole scheme again to take advantage of this new information. On the contrary, the online nature of the SGDW method allows one to refine the already computed estimator by only performing k' new steps of the algorithm.

5. NUMERICAL EXPERIMENTS

Before presenting the experimental validation of the proposed methods we give a brief presentation of the scatter-location family of distributions. Our experiments consider this family because the optimal transport maps between two laws in the scatter-location family can be described explicitly. This property facilitates the numerical computation/approximation of barycenters, as the various iterative algorithms described so far take a more amenable form. See [7] for further examples.

5.1. Location-Scatter family. We follow the setting of [4]: Given a fixed distribution $\tilde{m} \in \mathcal{W}_{2,ac}(\mathbb{R}^q)$, referred to as *generator*, the associated location-scatter family is given by

$$\mathcal{F}(\tilde{m}) := \{\mathcal{L}(A\tilde{x} + b) \mid A \in \mathcal{M}_+^{q \times q}, b \in \mathbb{R}^q, \tilde{x} \sim \tilde{m}\},$$

where $\mathcal{M}_+^{q \times q}$ is the set of symmetric positive definite matrices of size $q \times q$. Without loss of generality we can assume that \tilde{m} has zero mean and identity covariance. Note that $\mathcal{F}(\tilde{m})$ is the multivariate normal family if \tilde{m} is the standard multivariate normal distribution.

The optimal map between two members $m_1 = \mathcal{L}(A_1\tilde{x} + b_1)$ and $m_2 = \mathcal{L}(A_2\tilde{x} + b_2)$ of $\mathcal{F}(\tilde{m})$ is explicit, given by $T_{m_1}^{m_2}(x) = A(x - b_1) + b_2$ where $A = A_1^{-1}(A_1A_2^2A_1)^{1/2}A_1^{-1} \in \mathcal{M}_+^{q \times q}$. This family of optimal maps contains the identity and is closed under convex combination.

If Γ is supported on $\mathcal{F}(\tilde{m})$, then its 2-Wasserstein barycenter \hat{m} belongs to $\mathcal{F}(\tilde{m})$. In fact, call its mean \hat{b} and its covariance matrix $\hat{\Sigma}$. Since the optimal map from \hat{m} to m is $T_{\hat{m}}^m(x) = A_{\hat{m}}^m(x - \hat{b}) + b_m$, where $A_{\hat{m}}^m = \hat{\Sigma}^{-1/2}(\hat{\Sigma}^{1/2}\Sigma_m\hat{\Sigma}^{1/2})^{1/2}\hat{\Sigma}^{-1/2}$, and we know that $\int T_{\hat{m}}^m(x)\Gamma(dm) = x$, \hat{m} -almost surely, then we must have that $\int A_{\hat{m}}^m\Gamma(dm) = I$, since clearly $\hat{b} = \int b_m\Gamma(dm)$, and as a consequence $\hat{\Sigma} = \int (\hat{\Sigma}^{1/2}\Sigma_m\hat{\Sigma}^{1/2})^{1/2}\Gamma(dm)$.

A stochastic gradient descent iteration, starting from a distribution $\mu = \mathcal{L}(A_0\tilde{x} + b_0)$, sampling some $m = \mathcal{L}(A_m\tilde{x} + b_m) \sim \Gamma$, and with step γ , produces the measure $\nu = T_0^{\gamma,m}(\mu) := ((1 - \gamma)I + \gamma T_{\mu}^m)(\mu)$. If \tilde{x} has a multivariate distribution $\tilde{F}(x)$, then μ has distribution $F_0(x) = \tilde{F}(A_0^{-1}(x - b_0))$ with mean b_0 and covariance $\Sigma_0 = A_0^2$. We have that $T_0^{\gamma,m}(x) = ((1 - \gamma)I + \gamma A_{\mu}^m)(x - b_0) + \gamma b_m + (1 - \gamma)b_0$ with $A_{\mu}^m := A_0^{-1}(A_0A_m^2A_0)^{1/2}A_0^{-1}$. Then ν has distribution

$$F_1(x) = F_0([T_0^{\gamma,m}]^{-1}(x)) = \tilde{F}([(1 - \gamma)A_0 + \gamma A_{\mu}^m A_0]^{-1}(x - \gamma b_m - (1 - \gamma)b_0)),$$

with mean $b_1 = (1 - \gamma)b_0 + \gamma b_m$ and covariance

$$\begin{aligned} \Sigma_1 &= A_1^2 = [(1 - \gamma)A_0 + \gamma A_0^{-1}(A_0A_m^2A_0)^{1/2}][(1 - \gamma)A_0 + \gamma(A_0A_m^2A_0)^{1/2}A_0^{-1}] \\ &= A_0^{-1}[(1 - \gamma)A_0^2 + \gamma(A_0A_m^2A_0)^{1/2}][(1 - \gamma)A_0^2 + \gamma(A_0A_m^2A_0)^{1/2}]A_0^{-1} \\ &= A_0^{-1}[(1 - \gamma)A_0^2 + \gamma(A_0A_m^2A_0)^{1/2}]^2A_0^{-1}. \end{aligned}$$

The batch stochastic gradient descent iteration is characterized by

$$\begin{aligned} b_1 &= (1 - \gamma)b_0 + \frac{\gamma}{S} \sum_{i=1}^S b_{m_i} \\ A_1^2 &= A_0^{-1}[(1 - \gamma)A_0^2 + \frac{\gamma}{S} \sum_{i=1}^S (A_0A_{m_i}^2A_0)^{1/2}]^2A_0^{-1}. \end{aligned}$$

5.2. Experiment. We considered a model within a location-scatter family (LS), with generator \tilde{m} on \mathbb{R}^{15} with independent coordinates, as follows:

- coordinates 1 to 5 are standard Normal distributions
- coordinates 6 to 10 are standard Laplace distributions, and
- coordinates 11 to 15 are standard Student's t -distributions (3 degrees of freedom).

Fig. 2 shows samples (uni- and bi-variate marginals) from coordinates $\{1, 2, 6, 7, 11, 12\}$ of \tilde{m} .

In this LS family, we chose the true model m_0 with location vector $b \in \mathbb{R}^{15}$ defined as $b_i = i - 1$ for $i = 1, \dots, 15$, and scatter matrix $A = \Sigma^{1/2}$ with $\Sigma_{i,j} = K\left(\left(\frac{i-1}{14}\right)^{1.1}, \left(\frac{j-1}{14}\right)^{1.1}\right)$ for $i, j = 1, \dots, 15$ ², with the covariance (kernel) function $K(i, j) = \varepsilon\delta_{ij} + \sigma \cos(\omega(i - j))$. Given the parameters ε, σ and ω , the so constructed covariance matrix will be denoted $\Sigma_{\varepsilon, \sigma, \omega}$. We chose the parameters $\varepsilon = 0.01$, $\sigma = 1$ and $\omega = 5.652 \approx 1.8\pi$ for m_0 . Thus, under the true model m_0 the coordinates can be negatively/positively correlated and there is also a coordinate-independent noise component due to the Kronecker delta δ_{ij} . Fig. 3 shows the covariance matrix and three coordinates of the *true* model m_0 .

The model prior Π is the push-forward induced by a prior over the mean vector b and the parameters of the covariance $\Sigma_{\varepsilon, \sigma, \omega}$, chosen independently according to :

$$p(b, \Sigma_{\varepsilon, \sigma, \omega}) = \mathcal{N}(b|0, I) \text{Exp}(\varepsilon|20) \text{Exp}(\sigma|1) \text{Exp}(\omega^{-1}|15), \quad (5.1)$$

where $\text{Exp}(\cdot|\lambda)$ is a exponential density with rate λ . Given n samples from the true model m_0 (also referred to as *observations* or *data points*), k samples are produced from the posterior measure Π_n using Markov chain Monte Carlo (MCMC).

The numerical analysis presented in what follows focuses on the behavior of the BWB as a function of both the number of samples k and the number of data points n .

²We chose $\left(\frac{i-1}{14}\right)^{1.1}$ for $j = 1, \dots, 15$ because this defines a non-uniform grid over $[0, 1]$.

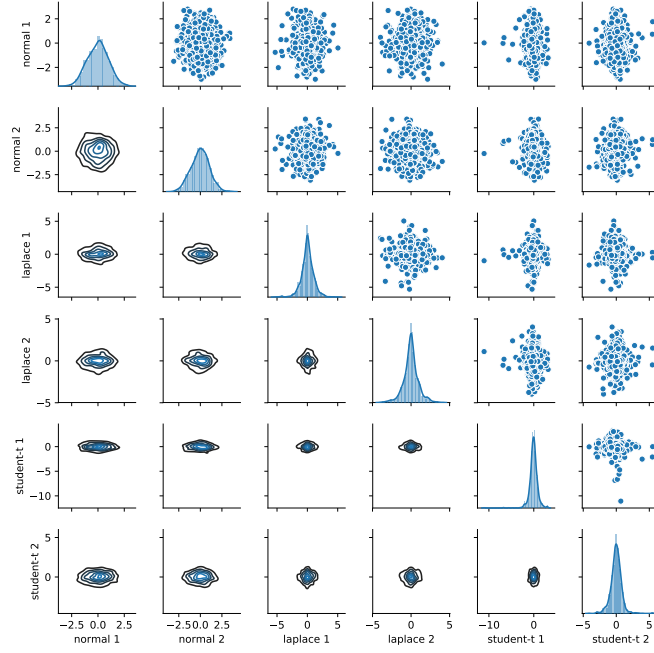


FIGURE 2. Samples from the univariate (diagonal) and bivariate (off-diagonal) marginals for 6 coordinates from the generator distribution \tilde{m} . The diagonal and lower triangular plots are smoothed histograms, whereas the upper-diagonal ones are collections of samples.

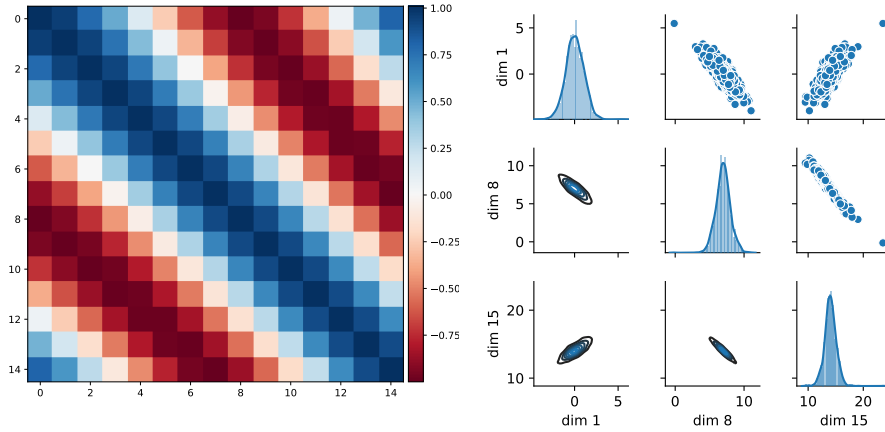


FIGURE 3. True model m_0 : covariance matrix (left), and univariate and bivariate marginals for dimensions 1, 8 and 15 (right). Notice that some coordinates are positively or negatively correlated, and some may even be close to uncorrelated.

Numerical consistency of the empirical posterior. We first validate the empirical measure $\Pi_n^{(k)}$ as a consistent sample version of the true posterior under the W_2 distance, that is, we check that $W_2(\Pi_n^{(k)}, \delta_{m_0}) \rightarrow W_2(\Pi_n, \delta_{m_0})$ for large k . We estimate $W_2(\Pi_n^{(k)}, \delta_{m_0})$ 10 times for each combination of the (number of) observations n and samples k in the following sets

- $k \in \{1, 5, 10, 20, 50, 100, 200, 500, 1000\}$

- $n \in \{10, 20, 50, 100, 200, 500, 1000, 2000, 5000, 10000\}$

Fig. 4 shows the 10 estimates of $W_2(\Pi_n^{(k)}, \delta_{m_0})$ for different values of k (in the x -axis) and of n (color coded). Notice how the estimates become more concentrated for larger k and that the Wasserstein distance between the empirical measure $\Pi_n^{(k)}$ and the true model m_0 decreases for larger n . Additionally, Table 1 shows that the standard deviation of the 10 estimates of $W_2(\Pi_n^{(k)}, \delta_{m_0})$ decreases as either n or k increases.

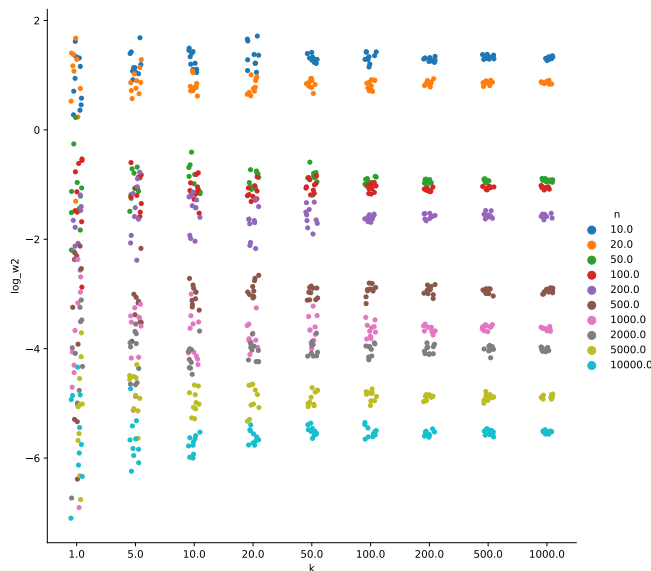


FIGURE 4. Wasserstein distance between the empirical measure $\Pi_n^{(k)}$ and δ_{m_0} in logarithmic scale for different number of observations n (color coded) and samples k (x -axis). For each pair (n, k) , 10 estimates of $W_2(\Pi_n^{(k)}, \delta_{m_0})$ are shown.

TABLE 1. Standard deviation of $W_2^2(\Pi_n^{(k)}, \delta_{m_0})$, using 10 simulations, for different values of observations n and samples k .

$n \setminus k$	1	5	10	20	50	100	200	500	1000
10	1.2506	0.8681	0.5880	0.9690	0.2354	0.3440	0.1253	0.1330	0.0972
20	1.5168	0.5691	0.3524	0.3182	0.1850	0.1841	0.1049	0.0811	0.0509
50	0.3479	0.0948	0.1275	0.0572	0.0623	0.0229	0.0157	0.0085	0.0092
100	0.2003	0.1092	0.0712	0.0469	0.0431	0.0254	0.0087	0.0079	0.0084
200	0.0749	0.1249	0.0717	0.0533	0.0393	0.0101	0.0092	0.0109	0.0072
500	0.0478	0.0285	0.0093	0.0086	0.0053	0.0056	0.0045	0.0023	0.0022
1000	0.0299	0.0113	0.0113	0.0064	0.0067	0.0036	0.0016	0.0012	0.0007
2000	0.0145	0.0071	0.0040	0.0031	0.0027	0.0019	0.0014	0.0011	0.0006
5000	0.0072	0.0031	0.0015	0.0018	0.0010	0.0007	0.0004	0.0005	0.0002
10000	0.0038	0.0020	0.0005	0.0005	0.0004	0.0004	0.0002	0.0002	0.0001

Wasserstein distance between the empirical barycenter and the true model. For each empirical posterior $\Pi_n^{(k)}$, we computed the empirical Wasserstein barycenter $\hat{m}_2^{(n,k)}$ as suggested in Remark 4.11. Thus, we used the iterative GDW procedure in eq. (4.2), namely the (deterministic) gradient descent method, and repeated this calculation 10 times. As a stopping criterion for gradient descent, we considered the relative variation of the W_2 cost; the computation was terminated when this cost fell below 10^{-4} . Fig. 5 shows the W_2 distances

between the so computed barycenters and the true model, while Table 2 shows the average distance for each pair (n, k) . Notice that, in general, both the average and standard deviation of the barycenters decrease as either n or k increases, yet for large values (e.g., $n = 2000, 5000$) numerical issues appear.

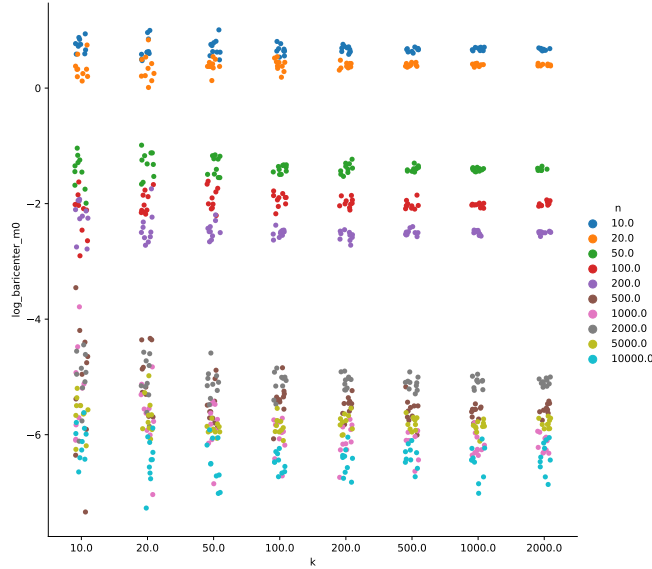


FIGURE 5. W_2 distance between the empirical barycenters $\hat{m}_2^{(n,k)}$ and the true model m_0 in logarithmic scale for different number of observations n (color coded) and samples k (x -axis). For each pair (n, k) , 10 estimates of $W_2(\hat{m}_2^{(n,k)}, m_0)$ are shown.

TABLE 2. Sample average of $W_2^2(\hat{m}_2^{(n,k)}, m_0)$, using 10 simulations, for different values of observations n and samples k .

n / k	10	20	50	100	200	500	1000	2000
10	2.1294	2.0139	2.0384	1.9396	1.9608	1.9411	1.9699	1.9548
20	1.4382	1.4498	1.4826	1.4973	1.4785	1.4953	1.4955	1.4914
50	0.2455	0.2759	0.2639	0.2468	0.2499	0.2483	0.2443	0.2454
100	0.1211	0.1387	0.1509	0.1458	0.1379	0.1328	0.1318	0.1349
200	0.1116	0.0922	0.0859	0.0817	0.0777	0.0824	0.0820	0.0819
500	0.0094	0.0077	0.0043	0.0047	0.0041	0.0038	0.0037	0.0039
1000	0.0068	0.0039	0.0031	0.0025	0.0023	0.0022	0.0021	0.0021
2000	0.0072	0.0066	0.0063	0.0062	0.0063	0.0060	0.0062	0.0062
5000	0.0037	0.0037	0.0028	0.0029	0.0031	0.0031	0.0028	0.0030
10000	0.0023	0.0017	0.0017	0.0015	0.0016	0.0017	0.0016	0.0017

Distance between the empirical barycenter and the Bayesian model average. We then compared the empirical Wasserstein barycenters $\hat{m}_2^{(n,k)}$ to the standard Bayesian model averages, denoted here $\bar{m}^{(n,k)}$, in terms of their distance to the true model m_0 , for $n = 1000$ observations. To that end, we estimated the W_2 distances via empirical approximations with 1000 samples for each model based on [21]. We simulated this procedure 10 times for $k \in \{10, 20, 50, 100, 200, 500, 1000\}$. Fig. 6 shows the sample average and variance of the W_2 distances of the Wasserstein barycenters and Bayesian model averages. The empirical barycenter is seen to be closer to the true model than the model average, regardless of the number of MCMC samples k .

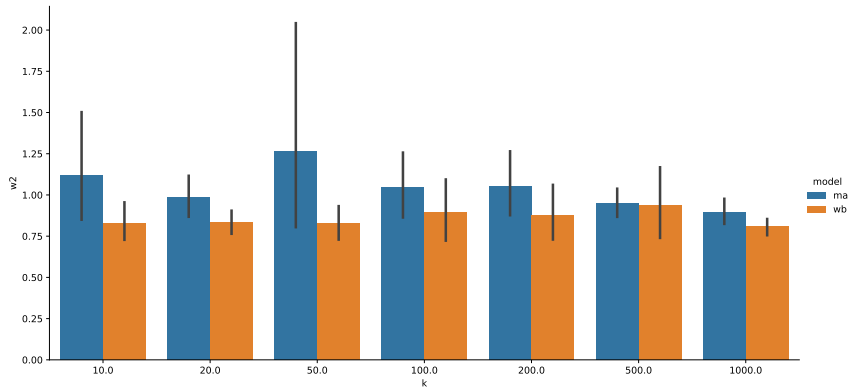


FIGURE 6. Averages (bars) and standard deviations (vertical lines) of $W_2^2(\hat{m}_n^{(k)}, m_0)$ denoted as **wb** in orange, and $W_2^2(\bar{m}_n^{(k)}, m_0)$ denoted as **ma** in blue, for $n = 1000$ and different numbers of samples k . We considered 10 simulations for each k .

Computation of the Wasserstein barycenter with batch-SGDW. Lastly, we compared the empirical barycenters $\hat{m}_2^{(m,k)}$ to the barycenter obtained by the batch SGDW method in eq. (4.8) with different batch sizes. Here, we shall denote the latter by $\hat{m}_2^{(n,t,s)}$ with n the number of observations, t the steps of the algorithm, and s the batch size. Fig. 7 shows the evolution of the W_2^2 distance between the stochastic gradient descent sequences and the true model m_0 for $n \in \{10, 20, 50, 100, 200, 500, 1000\}$ observations and batches of sizes $s \in \{1, 15\}$, with step-size $\gamma_t = \frac{1}{t}$ for $t = 1, \dots, 200$. Notice from Fig. 7 that the larger the batch, the more concentrated the trajectories of $\hat{m}_{n,s}$ become. Additionally, Table 3 summarizes the means of the distance W_2^2 to the true model m_0 , using the sequences after $t = 100$ against the empirical estimator using all the simulations with $k \geq 100$. Finally, Table 4 shows the standard deviation of the distance W_2^2 to the true model m_0 , which can be seen to decrease as the batch size grows. Critically, we observe that for batch sizes $s \geq 5$ the stochastic estimation was *better* than its empirical counterpart, i.e., it had smaller variance with similar (or even smaller) bias. This is noteworthy given the fact that computing our Wasserstein barycenter estimator via the batch stochastic gradient descent method is computationally less demanding than computing it via the empirical method.

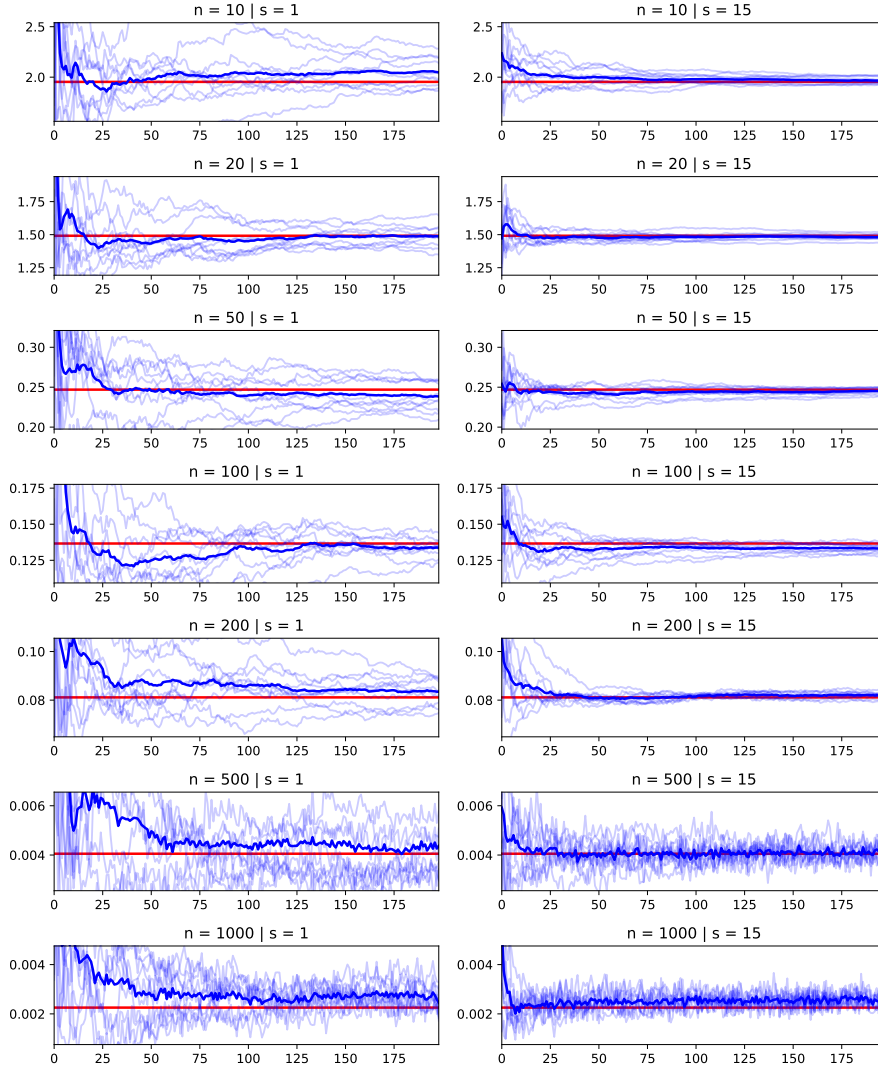


FIGURE 7. Evolution of the W_2^2 cost for 10 realizations of the SGD sequence computing the BWB and their mean (blue), versus an empirical barycenter estimator (red), for $n = 10, 20, 50, 100, 200, 500, 1000$ and batches sizes $s = 1, 15$.

TABLE 3. Means of W_2^2 of the stochastic gradient estimations (using the sequences with $t \geq 100$) and that of the empirical estimator (using the simulations with $k \geq 100$), across different combinations of observations n and batch size s .

n/s	1	2	5	10	15	20	empirical
10	2.0421	2.0091	1.9549	1.9721	1.9732	1.9712	1.9532
20	1.4819	1.4868	1.5100	1.4852	1.4840	1.4891	1.4916
50	0.2406	0.2512	0.2465	0.2427	0.2444	0.2460	0.2469
100	0.1340	0.1392	0.1340	0.1349	0.1334	0.1338	0.1366
200	0.0843	0.0811	0.0819	0.0807	0.0820	0.0819	0.0811
500	0.0044	0.0042	0.0039	0.0039	0.0041	0.0040	0.0041

TABLE 4. Std. deviations of W_2^2 of the stochastic gradient estimations (using the sequences with $t \geq 100$) and that of empirical estimator (using the simulations with $k \geq 100$), across different combinations of observations n and batch size s .

n / s	1	2	5	10	15	20	empirical
10	0.1836	0.1071	0.0526	0.0474	0.0397	0.0232	0.0916
20	0.0751	0.0565	0.0553	0.0189	0.0253	0.0186	0.0790
50	0.0210	0.0174	0.0072	0.0084	0.0050	0.0039	0.0138
100	0.0102	0.0076	0.0049	0.0048	0.0035	0.0023	0.0112
200	0.0074	0.0045	0.0021	0.0035	0.0013	0.0017	0.0047
500	0.0016	0.0007	0.0005	0.0004	0.0004	0.0004	0.0009
1000	0.0005	0.0006	0.0004	0.0004	0.0003	0.0003	0.0005

Conclusion. Based on this illustrative numerical example, we can conclude:

- the empirical posterior constructed using MCMC is consistent under the W_2 distance and therefore it can be relied upon to compute Wasserstein barycenters,
- the empirical Wasserstein barycenter estimator tends to converge faster (and with lower variance) to the true model than the empirical Bayesian model average,
- computing the population Wasserstein barycenter estimator via batch stochastic gradient descent is promising as an alternative to computing the empirical barycenter (i.e., to applying the deterministic gradient descent method to a finitely-sampled posterior).

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APPENDIX A. BAYES ESTIMATORS AS GENERALIZED MODEL AVERAGES

We prove here Proposition 1.5. For notational simplicity we omit the subscripts of estimators and normalizing constants.

i) Squared L_2 -distance: By Fubini's theorem, minimizing $R_L(\bar{m}|D)$ in this case amounts to minimize

$$\bar{m} \mapsto \frac{1}{2} \int_{\mathcal{X}} \left\{ \int_{\mathcal{M}} (m(x) - \bar{m}(x))^2 \Pi_n(dm) \right\} \lambda(dx),$$

over the set of densities. But the optimal value cannot be better than if we minimize pointwise the term $\bar{m}(x)$ inside the curly brackets, obtaining the candidate

$$\hat{m}_{BMA}(x) = \int_{\mathcal{M}} m(x) \Pi_n(dm) = \mathbb{E}[m](x).$$

As this pointwise minimizer is already a probability density, we conclude.

ii) Squared Hellinger distance: Writing

$$H^2(m, \bar{m}) = \frac{1}{2} \int_{\mathcal{X}} \left(\sqrt{m(x)} - \sqrt{\bar{m}(x)} \right)^2 \lambda(dx) = 1 - \int_{\mathcal{X}} \sqrt{m(x)\bar{m}(x)} \lambda(dx),$$

we see that optimizing the associated Bayes risk amounts to maximizing over $\bar{m} \in L^1(\mathcal{X}, \lambda)$ the concave functional $\bar{m} \mapsto \int_{\mathcal{X}} \sqrt{\bar{m}(x)} f(x) \lambda(dx)$ with $\sqrt{y} = -\infty$ for $y < 0$ and $f(x) = \int_{\mathcal{M}} \sqrt{m(x)} \Pi_n(dm)$, under the constraint $\int_{\mathcal{X}} \bar{m}(x) \lambda(dx) = 1$. Notice by the Cauchy-Schwarz inequality that this functional is finite if (and only if) $\bar{m} \geq 0$, λ -a.e. Hence, introducing γ a real Lagrange multiplier for the constraint, we need to find a critical point $\bar{m} \geq 0$ of the concave functional

$$\bar{m} \mapsto \int_{\mathcal{X}} \sqrt{\bar{m}(x)} f(x) \lambda(dx) + \gamma \left(\int_{\mathcal{X}} \bar{m}(x) \lambda(dx) - 1 \right),$$

which must thus be a maximum. Again, as in i), we cannot do better in this case than if we maximize for each $x \in \mathcal{X}$ the concave functional $y \mapsto \sqrt{y} f(x) + \gamma y$ over $y \geq 0$. Finding for each x the critical point y in terms of γ and integrating then w.r.t. $\lambda(dx)$ to get rid of γ , we find the extremum of $R_{H^2}(\bar{m}|D)$ is attained when \bar{m} equals the Bayesian *square* model average:

$$\hat{m}(x) = \frac{1}{Z} \left(\int_{\mathcal{M}} \sqrt{m(x)} \Pi_n(dm) \right)^2 \text{ with} \\ Z = \int_{\mathcal{X}} \left(\int_{\mathcal{M}} \sqrt{m(x)} \Pi_n(dm) \right)^2 \lambda(dx).$$

iii) Forward Kullback-Leibler divergence: the loss function $L(m, \bar{m})$ is now

$$D_{KL}(m||\bar{m}) = \int_{\mathcal{X}} m(x) \ln \frac{m(x)}{\bar{m}(x)} \lambda(dx),$$

and so the associated Bayes risk can be written as

$$\begin{aligned} R_{D_{KL}}(\bar{m}|D) &= \int_{\mathcal{M}} \int_{\mathcal{X}} m(x) \ln \frac{m(x)}{\bar{m}(x)} \lambda(dx) \Pi_n(dm) \\ &= \int_{\mathcal{X}} \int_{\mathcal{M}} m(x) \ln m(x) \Pi_n(dm) \lambda(dx) - \int_{\mathcal{X}} \int_{\mathcal{M}} m(x) \Pi_n(dm) \ln \bar{m}(x) \lambda(dx) \\ &= C - \int_{\mathcal{X}} \mathbb{E}[m](x) \ln \bar{m}(x) \lambda(dx). \end{aligned}$$

Introducing the Boltzmann entropy of $\mathbb{E}[m]$ and adjusting the constant C we get that

$$\begin{aligned} R_{D_{KL}}(\bar{m}|D) &= C' + \int_{\mathcal{X}} \mathbb{E}[m](x) \ln \mathbb{E}[m](x) \lambda(dx) - \int_{\mathcal{X}} \mathbb{E}[m](x) \ln \bar{m}(x) \lambda(dx) \\ &= C' + D_{KL}(\mathbb{E}[m]||\bar{m}), \end{aligned}$$

so the optimizer of $R_{D_{KL}}(\bar{m}|D)$ is the Bayesian model average.

iv) Reverse Kullback-Leibler divergence: since $D_{RKL}(m|\bar{m}) = D_{KL}(\bar{m}|m)$, we have

$$\begin{aligned} R_{D_{RKL}}(\bar{m}|D) &= \int_{\mathcal{M}} \int_{\mathcal{X}} \bar{m}(x) \ln \frac{\bar{m}(x)}{m(x)} \lambda(dx) \Pi_n(dm) \\ &= \int_{\mathcal{X}} \bar{m}(x) \ln \bar{m}(x) \lambda(dx) - \int_{\mathcal{X}} \bar{m}(x) \int_{\mathcal{M}} \ln m(x) \Pi_n(dm) \lambda(dx) \\ &= \int_{\mathcal{X}} \bar{m}(x) \ln \bar{m}(x) \lambda(dx) - \int_{\mathcal{X}} \bar{m}(x) \ln \exp \mathbb{E}[\ln m(x)] \lambda(dx) \\ &= \int_{\mathcal{X}} \bar{m}(x) \ln \frac{\bar{m}(x)}{\exp \mathbb{E}[\ln m(x)]} \lambda(dx). \end{aligned}$$

Denote by Z the normalization constant so that $\frac{1}{Z} \int_{\mathcal{X}} \exp \mathbb{E}[\ln m(x)] \lambda(dx) = 1$. Then,

$$\begin{aligned} R_{D_{RKL}}(\bar{m}|D) + \ln Z &= \int_{\mathcal{X}} \bar{m}(x) \ln \frac{\bar{m}(x)}{\exp \mathbb{E}[\ln m(x)]} \lambda(dx) + \int_{\mathcal{X}} \bar{m}(x) \ln Z \lambda(dx) \\ &= \int_{\mathcal{X}} \bar{m}(x) \ln \frac{\bar{m}(x)}{\frac{1}{Z} \exp \mathbb{E}[\ln m(x)]} \lambda(dx) \\ &= D_{RKL}\left(\frac{1}{Z} \exp \mathbb{E}[\ln m(x)] \middle| \bar{m}\right). \end{aligned}$$

Therefore, the extremum of $R_{D_{RKL}}(\bar{m}|D)$ is attained when the last expression vanishes, in other words when \bar{m} is the Bayesian *exponential* model average given by

$$\hat{m}(x) = \frac{1}{Z} \exp \int_{\mathcal{M}} \ln m(x) \Pi_n(dm).$$

APPENDIX B. WASSERSTEIN BARYCENTERS

Proof of Theorem 2.7. Assume $\Gamma \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$. Then $B_V := \inf_{\nu \in \mathcal{M}} V_p(\nu)$ is finite. Now, let $\{\nu_n\} \subseteq \mathcal{M}$ be such that

$$\int_{\mathcal{W}_p(\mathcal{X})} W_p(\nu_n, m)^p \Gamma(dm) \searrow B_V \text{ as } n \rightarrow \infty.$$

For n large enough we have

$$W_p\left(\nu_n, \int_{\mathcal{W}_p(\mathcal{X})} m \Gamma(dm)\right)^p \leq \int_{\mathcal{W}_p(\mathcal{X})} W_p(\nu_n, m)^p \Gamma(dm) \leq B_V + 1 =: K,$$

by convexity of optimal transport costs. From this we derive that (for every x)

$$\sup_n \int_{\mathcal{X}} d(x, y)^p \nu_n(dy) < \infty.$$

By Markov inequality this shows, for each $\varepsilon > 0$, that there is ℓ large enough such that $\sup_n \nu_n(\{y \in \mathcal{X} : d(x, y) > \ell\}) \leq \varepsilon$. As explained in [33], the assumptions on \mathcal{X} imply that $\{y \in \mathcal{X} : d(x, y) \leq \ell\}$ is compact (Hopf-Rinow theorem), and so we deduce the tightness of $\{\nu_n\}$. By Prokhorov theorem, up to selection of a subsequence, there exists $\nu \in \mathcal{M}$ which is its weak limit. By Fatou's lemma:

$$B_V = \lim \int_{\mathcal{W}_p(\mathcal{X})} W_p(\nu_n, m)^p \Gamma(dm) \geq \int \liminf W_p(\nu_n, m)^p \Gamma(dm) \geq \int W_p(\nu, m)^p \Gamma(dm),$$

hence ν is a p -Wasserstein barycenter. For the converse implication, see Remark 2.6. \square

APPENDIX C. A CONDITION FOR EXISTENCE OF BARYCENTERS OF BAYESIAN POSTERiors

We last provide a general condition on the prior Π ensuring that, for given $p \geq 1$,

$$\Pi_n \in \mathcal{W}_p(\mathcal{W}_p(\mathcal{X})) \text{ for all possible data points } (x_1, \dots, x_n) \in \mathcal{X}^n \text{ and all } n.$$

Definition C.1. We say that $\Pi \in \mathcal{P}(\mathcal{P}(\mathcal{X}))$ is p -integrable after updates if it satisfies the conditions

(1) For all $x \in \mathcal{X}, \ell > 1$:

$$\int_{\mathcal{M}} m(x)^\ell \Pi(dm) < \infty.$$

(2) For some $y \in \mathcal{X}$ and $\varepsilon > 0$:

$$\int_{\mathcal{M}} \left(\int_{\mathcal{X}} d(y, z)^p m(dz) \right)^{1+\varepsilon} \Pi(dm) < \infty.$$

Remark C.2. Condition (2) above can be intuitively denoted as $\Pi \in \mathcal{W}_{p+}(\mathcal{W}_p(\mathcal{X}))$. Of course, one has $\mathcal{W}_{p+}(\mathcal{W}_p(\mathcal{X})) \subseteq \mathcal{W}_p(\mathcal{W}_p(\mathcal{X}))$.

Remark C.3. If $\Pi \in \mathcal{P}(\mathcal{W}_{p,ac}(\mathcal{X}))$ has finite support, then Conditions (1) and (2) are satisfied. On the other hand, if Π is supported on a scatter-location family (see Section 5.1) containing one element with a bounded density and a finite p -moment, then Conditions (1) and (2) are fulfilled if for example $\text{supp}(\Pi)$ is tight. Conditions (1) and (2) are also satisfied in Example 3.2.

Lemma C.4. *Suppose that Π is p -integrable after updates. Then, for each $x \in \mathcal{X}$, the measure*

$$\tilde{\Pi}(dm) := \frac{m(x)\Pi(dm)}{\int_{\mathcal{M}} \bar{m}(x)\Pi(d\bar{m})},$$

is also p -integrable after updates.

Proof. We verify Property (1) first. Let $\ell > 1$ and $\bar{x} \in \mathcal{X}$ given. Then

$$\int_{\mathcal{M}} m(\bar{x})^\ell m(x)\Pi(dm) \leq \left(\int_{\mathcal{M}} m(x)^s \Pi(dm) \right)^{1/s} \left(\int_{\mathcal{M}} m(\bar{x})^{t\ell} \Pi(dm) \right)^{1/t},$$

with s, t conjugate Hölder exponents. This is finite since Π fulfils Property (1).

We now establish Property (2). Let $y \in \mathcal{X}, \varepsilon > 0$. Then

$$\begin{aligned} & \int_{\mathcal{M}} \left(\int_{\mathcal{X}} d(y, z)^p m(dz) \right)^{1+\varepsilon} m(x)\Pi(dm) \\ & \leq \left(\int_{\mathcal{M}} m(x)^s \Pi(dm) \right)^{1/s} \left(\int_{\mathcal{M}} \left(\int_{\mathcal{X}} d(y, z)^p m(dz) \right)^{(1+\varepsilon)t} \Pi(dm) \right)^{1/t}. \end{aligned}$$

The first term in the r.h.s. is finite by Property (1). The second term in the r.h.s. is finite by Property (2), if we take ε small enough and t close enough to 1. We conclude. \square

Lemma C.5. *Suppose that Π is p -integrable after updates. Then for all $n \in \mathbb{N}$ and $(x_1, \dots, x_n) \in \mathcal{X}^n$, the posterior Π_n is also p -integrable after updates.*

Proof. By Lemma C.4, we obtain that Π_1 is integrable after updates. By induction, suppose Π_{n-1} has this property. Then as

$$\Pi_n(dm) = \frac{m(x_n)\Pi_{n-1}(dm)}{\int_{\mathcal{M}} \bar{m}(x_n)\Pi_{n-1}(d\bar{m})},$$

we likewise conclude that Π_n is p -integrable after updates. \square

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