Topics in Statistical Theory — Summary

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November 16, 2020

Contents

1 Basic concepts

1.1 Parametric vs nonparametric models

Definition 1.1. A *statistical model* is a family of possible data-generating mechanisms. If the parameter space Θ is finite-dimensional, we speak of a *parametric model*.

A model is called *well-specified* if there is a $\vartheta_0 \in \Theta$ for which the data was generated from the distribution with parameter ϑ_0, and otherwise it is called *misspecified*.

Recap 1.2. Let (Y_n) be a sequence of random vectors and Y a random vector.

1. We say that (Y_n) *converges almost surely* to Y, notation $Y_n \overset{\text{a.s.}}{\to} Y$, if $\mathbb{P}(Y_n \to Y) = 1$.
2. We say that (Y_n) *converges in probability* to Y, notation $Y_n \overset{\text{p}}{\to} Y$, if for every $\varepsilon > 0$ we have $\mathbb{P}(\|Y_n - Y\| > \varepsilon) \to 0$.
3. We say that (Y_n) *converges in distribution* to Y, notation $Y_n \overset{\text{d}}{\to} Y$, if $\mathbb{P}(Y_n \le y) \to \mathbb{P}(Y \le y)$ for all y where the distribution function of Y is continuous.

   This is equivalent to the condition that $\mathbb{E}[f(Y_n)] \to \mathbb{E}[f(Y)]$ for all bounded Lipschitz functions f.

It is known that $Y_n \overset{\text{a.s.}}{\to} Y \implies Y_n \overset{\text{p}}{\to} Y \implies Y_n \overset{\text{d}}{\to} Y$.

If (Y_n) is a sequence of random vectors and (a_n) is a positive sequence, then we write $Y_n = O_p(a_n)$ if, for all $\varepsilon > 0$, there exists $C > 0$ such that for sufficiently large n we have

$$\mathbb{P}\left(\frac{\|Y_n\|}{a_n} > C\right) < \varepsilon.$$

We write $Y_n = o_n(a_n)$ if $Y_n/a_n \overset{\text{p}}{\to} 0$.

In a well-specified parametric model, the maximum likelihood estimator (MLE) $\hat{\vartheta}_n$ typically satisfies $\hat{\vartheta}_n - \vartheta_0 \in O_p(n^{-1/2})$. On the other hand, if the model is misspecified, any inference can give very misleading results. To circumvent this problem, we consider *nonparametric models*, which make much weaker assumptions. Such infinite-dimensional models are much less vulnerable to model misspecification, however we will typically pay a price in terms of a slower convergence rate than in well-specified parametric models.

Example 1.3. Examples of nonparametric models include:

1. Assume $X_1, \ldots, X_n \overset{\text{iid}}{\sim} F$ for some unknown distribution function F.
2. Assume $X_1, \ldots, X_n \overset{\text{iid}}{\sim} f$ for some unknown density f belonging to a smoothness class.
3. Assume $Y_i = m(x_i) + \varepsilon_i$ $(i = 1, \ldots, n)$, where the x_i are known, m is unknown and belongs to some smoothness class, and the ε_i are i.i.d. with $\mathbb{E}(\varepsilon_i) = 0$ and $\mathrm{Var}(\varepsilon_i) = \sigma^2$.

1.2 Estimating an arbitrary distribution function

Definition 1.4. Let $\mathcal{F}$ denote the class of all distribution functions on $\mathbb{R}$ and suppose $X_1, \ldots, X_n \overset{\text{iid}}{\sim} F \in \mathcal{F}$. The *empirical distribution function* $\hat{F}_n$ of $X_1, \ldots, X_n$ is defined as

$$\hat{F}_n(x) := \frac{1}{n} \sum_{i=1}^{n} \mathbb{1}_{\{X_i \le x\}}.$$

Recap 1.5. The *strong law of large numbers* tells us that if (Y_n) are i.i.d. with finite mean μ, then $\bar{Y} \coloneqq \frac{1}{n}\sum_{i=1}^n Y_i \overset{\text{a.s.}}{\to} \mu$.

Note that the strong law of large numbers immediately implies that $\hat{F_n}(x)$ converges almost surely to $F(x)$ as $n \to \infty$. However, the following stronger result states that this convergence holds uniformly in x:

Theorem 1.6 (Glivenko-Cantelli)**.** *Let $X_1, X_2, \ldots \overset{\text{iid}}{\sim} F$. Then we have*

$$\sup_{x\in\mathbb{R}} \left|\hat{F_n}(x) - F(x)\right| \overset{\text{a.s.}}{\to} 0.$$

Proof. See lecture notes. The main idea of the proof is to "control" $\hat{F_n}$ in a finite number of points $x_1, \ldots, x_k$, and then deduce what happens between those points using the fact that distributions are increasing and right-continuous. On Wikipedia, a simplified proof can be found assuming that F is continuous, which still encapsulates the main idea. □

Theorem 1.7 (Dvoretzky-Kiefer-Wolfowitz)**.** *Under the conditions of theorem 1.6, for every $\varepsilon > 0$ it holds that*

$$\mathbb{P}_F\left(\sup_{x\in\mathbb{R}} \left|\hat{F_n}(x) - F(x)\right| > \varepsilon\right) \leq 2e^{-2n\varepsilon^2},$$

and this is a tight bound.

We will not prover this theorem, however, we will explore a few consequences. One of these consequences is the following:

Corollary 1.8 (Uniform Glivenko-Cantelli theorem)**.** *Under the conditions of theorem 1.6, for every $\varepsilon > 0$, it holds that*

$$\sup_{F\in\mathcal{F}} \mathbb{P}_F\left(\sup_{m\geq n}\sup_{x\in\mathbb{R}} \left|\hat{F_m}(x) - F(x)\right| > \varepsilon\right) \to 0 \quad \text{as } n \to \infty.$$

Proof. By a union bound, the DKW inequality, and convergence of the geometric series we have

$$\begin{aligned}\sup_{F\in\mathcal{F}} \mathbb{P}_F\left(\sup_{m\geq n}\sup_{x\in\mathbb{R}} \left|\hat{F_m}(x) - F(x)\right| > \varepsilon\right) &\leq \sup_{F\in\mathcal{F}} \sum_{m=n} \mathbb{P}_F\left(\sup_{x\in\mathbb{R}} \left|\hat{F_n}(x) - F(x)\right| > \varepsilon\right) \\ &\leq 2\sum_{m=n}^{\infty} e^{-2m\varepsilon^2},\end{aligned}$$

which converges to 0 as it is the tail of a converging sum. □

For another consequence, we consider the problem of finding a confidence band for F. Given $\alpha \in (0, 1)$, set $\varepsilon_n \coloneqq \sqrt{-\frac{1}{2n}\log(\alpha/2)}$. Then the DKW inequality tells us that

$$\mathbb{P}_F\left(\sup_{x\in\mathbb{R}} \left|\hat{F_n}(x) - F(x)\right| > \varepsilon_n\right) \leq \alpha,$$

or equivalently, that

$$\mathbb{P}_F\left(\hat{F_n}(x) - \varepsilon_n \leq F(x) \leq \hat{F_n}(x) + \varepsilon_n \text{ for all } x \in \mathbb{R}\right) \geq 1 - \alpha.$$

We can say even more.

Recap 1.9. For any distribution function F, its *quantile function* is defined as

$$F^{-1}\colon (0,1] \to \mathbb{R} \cup \{\infty\}\colon p \mapsto \inf\{x \in \mathbb{R} \mid F(x) \geq p\}.$$

When necessary, we also define $F^{-1}(0) \coloneqq \sup\{x \in \mathbb{R} \mid F(x) = 0\}$.

If $U \sim U(0,1)$ and $X \sim F$, then for any $x \in \mathbb{R}$ we have

$$\mathbb{P}(F^{-1}(U) \leq x) = \mathbb{P}(U \leq F(x)) = F(x) = \mathbb{P}(X \leq x).$$

This can be written simply as $F^{-1}(U) \overset{\mathrm{d}}{=} X$.

Let $U_1, \ldots, U_n \overset{\mathrm{iid}}{\sim} U(0,1)$ with empirical distribution function $\hat{G_n}$, and let $X_1, \ldots, X_n \overset{\mathrm{iid}}{\sim} F$. Then, we have

$$\hat{G_n}(F(x)) = \frac{1}{n}\sum_{i=1}^{n} \mathbb{1}_{\{U_i \leq F(x)\}} \overset{\mathrm{d}}{=} \frac{1}{n}\sum_{i=1}^{n} \mathbb{1}_{\{X \leq x\}} = \hat{F_n}(x),$$

where $\overset{\mathrm{d}}{=}$ means equality in distribution. It follows that

$$\sup_{x\in\mathbb{R}} \left|\hat{F_n}(x) - F(x)\right| \overset{\mathrm{d}}{=} \sup_{x\in\mathbb{R}} \left|\hat{G_n}(F(x)) - F(x)\right| \leq \sup_{t\in[0,1]} \left|\hat{G_n}(t) - t\right|,$$

with equality if F is continuous. We conclude that if F is continuous, the distribution of $\sup_{x\in\mathbb{R}} \left|\hat{F_n}(x) - F(x)\right|$ does not depend on F.

Other generalisations of theorem 1.6 include Uniform Laws of Large Numbers. Let $X, X_1, \ldots, X_n$ be i.i.d. on a measurable space $(\mathcal{X}, \mathcal{A})$, and $\mathcal{G}$ a class of measurable functions on $\mathcal{X}$. We say that $\mathcal{G}$ satisfies a ULLN if

$$\sup_{g\in\mathcal{G}} \left|\frac{1}{n}\sum_{i=1}^{n} g(X_i) - \mathbb{E}[g(X)]\right| \xrightarrow{\mathrm{a.s.}} 0.$$

In theorem 1.6, we showed that $\mathcal{G} = \left\{\mathbb{1}_{\{\cdot \leq x\}} \mid x \in \mathbb{R}\right\}$ satisfies a ULLN.

Recap 1.10. We recall the central limit theorem: if $X_1, \ldots, X_n$ are i.i.d. random variables with mean μ and variance $\sigma^2 < \infty$, then $\sqrt{n}\big(\bar{X}_n - \mu\big) \xrightarrow{\mathrm{d}} N(0, \sigma^2)$.

Dividing by σ yields

$$\frac{\sqrt{n}\big(\bar{X}_n - \mu\big)}{\sigma} \xrightarrow{\mathrm{d}} N(0,1),$$

and multiplying both sides by n and writing $V_i = \sum_{j=1}^{i} X_j$ we obtain

$$\frac{V_i - \mathbb{E}V_i}{\sqrt{\operatorname{Var}(V_i)}} \xrightarrow{\mathrm{d}} N(0,1).$$

Another extension starts with the observation that $\sqrt{n}\Big(\hat{F_n}(x) - F(x)\Big) \xrightarrow{\mathrm{d}} N(0, \sigma^2)$, where

$$\sigma^2 = \operatorname{Var}(\mathbb{1}_{\{X\leq x\}}) = \mathbb{E}[\mathbb{1}_{X\leq x}^2] - \mathbb{E}[\mathbb{1}_{X\leq x}]^2 = F(x) - F(x)^2 = F(x)(1 - F(x)).$$

This can be strengthened by considering $(\sqrt{n}\Big(\hat{F_n}(x) - F(x)\Big) \xrightarrow{\mathrm{d}} N(0, \sigma^2) \mid x \in \mathbb{R})$ as a stochastic process.

1.3 Order statistics and quantiles

Definition 1.11. Let $X_1, \dots, X_n \overset{\text{iid}}{\sim} F \in \mathcal{F}$. The *order statistics* are the ordered samples $X_{(1)} \leq \dots \leq X_{(n)}$ (where the original order is preserved in case of a tie).

The order statistics of the uniform distribution can be computed explicitly:

Proposition 1.12. *Let $U_1, \dots, U_n \overset{\text{iid}}{\sim} U(0,1)$, let $Y_1, \dots, Y_{n+1} \overset{\text{iid}}{\sim} \text{Exp}(1)$, and write $S_j \coloneqq \sum_{i=1}^j Y_j$ ($j = 1, \dots, n+1$). Then*

$$U_{(j)} \overset{\text{d}}{=} \frac{S_j}{S_{n+1}} \sim \text{Beta}(j, n-j+1) \quad \textit{for } j = 1, \dots, n.$$

Proof. See example sheet 1, question 1. □

Definition 1.13. Let $X_1, \dots, X_n \overset{\text{iid}}{\sim} F$. Then the *sample quantile function* is defined as

$$\hat{F}_n^{-1}(p) = \inf\left\{x \in \mathbb{R} \mid \hat{F}_n(x) \geq p\right\}.$$

Proposition 1.14. *It holds that $\hat{F}_n^{-1}(p) = X_{(\lceil np \rceil)}$.*

Proof. By definition, $\hat{F}_n^{-1}(p)$ is the smallest value of x for which $\hat{F}(x)$ is larger than p. Note that

$$\hat{F}(x) \geq p \iff \frac{1}{n}\sum_{i=1}^n \mathbb{1}_{\{X_i \leq x\}} \geq p \iff \sum_{i=1}^n \mathbb{1}_{\{X_i \leq x\}} \geq np \iff \sum_{i=1}^n \mathbb{1}_{\{X_i \leq x\}} \geq \lceil np \rceil.$$

The smallest value of x for which this occurs is the smallest value of x such that exactly $\lceil np \rceil$ of the variables $X_1, \dots, X_n$ satisfy $X_i \leq x$. We conclude that $\hat{F}_n^{-1}(p) = X_{(\lceil np \rceil)}$. □

For $p = \frac{1}{2}$ for example, this proposition tells us that $\hat{F}_n^{-1}(p) = X_{(\lceil n/2 \rceil)}$, the median of the data. We now explore the distribution of $X_{(\lceil np \rceil)}$.

Recap 1.15. We recall two theorems. The first is *Slutsky's theorem*:

Theorem 1.16. *Let (Y_n) and (Z_n) be sequences of random vectors with $Y_n \xrightarrow{\text{d}} Y$ and $Z_n \xrightarrow{\text{p}} c$ for some constant c. If g is a continuous real-valued function, then $g(Y_n, Z_n) \xrightarrow{\text{d}} g(Y, c)$.*

The second is the *delta method*:

Theorem 1.17. *Let (Y_n) be a sequence of random vectors such that $\sqrt{n}(Y_n - \mu) \xrightarrow{\text{d}} Z$. If $g \colon \mathbb{R}^d \to \mathbb{R}$ is differentiable at μ, then*

$$\sqrt{n}(g(Y_n) - g(\mu)) \xrightarrow{\text{d}} g'(\mu)Z.$$

Lemma 1.18. *If $U_1, \dots, U_n \overset{\text{iid}}{\sim} U(0,1)$ and $p \in (0,1)$, then $\sqrt{n}\big(U_{\lceil np \rceil} - p\big) \xrightarrow{\text{d}} N(0, p(1-p))$.*

Proof. Let $Y_1, \dots, Y_{n+1} \overset{\text{iid}}{\sim} \text{Exp}(1)$, $V_n \coloneqq \sum_{i=1}^{\lceil np \rceil} Y_i$ and $W_n \coloneqq \sum_{i=\lceil np \rceil + 1}^{n+1} Y_i$. Then V_n and W_n are independent, and we have seen that $U_{\lceil np \rceil} \sim \frac{V_n}{V_n + W_n}$.

Noting that $\mathbb{E}V_n = \text{Var}(V_n) = \lceil np \rceil$ we find

$$\begin{aligned}
\sqrt{n}\left(\frac{V_n}{n} - p\right) &= \frac{\sqrt{\lceil np \rceil}}{\sqrt{n}}\left(\frac{V_n - \lceil np \rceil}{\sqrt{\lceil np \rceil}}\right) + \frac{\lceil np \rceil - np}{\sqrt{n}} \\
&= \frac{\sqrt{\lceil np \rceil}}{\sqrt{n}}\left(\frac{V_n - \mathbb{E}V_n}{\sqrt{\text{Var}(V_n)}}\right) + \frac{\lceil np \rceil - np}{\sqrt{n}}.
\end{aligned}$$

Now, by the central limit theorem, the term between brackets converges to a standard $N(0,1)$ distribution. The term $\sqrt{\lceil np\rceil}\sqrt{n}$ converges to $\sqrt{p}$ and the term $(\lceil np\rceil - np)/\sqrt{n}$ converges to 0, so by Slutsky's lemma, we find

$$\sqrt{n}\left(\frac{V_n}{n} - p\right) \xrightarrow{\mathrm{d}} \sqrt{p}N(0,1) = N(0,p).$$

An analogous calculation shows that $\sqrt{n}\big(\frac{W_n}{n} - (1-p)\big) \to N(0,1-p)$.

Now we define $g\colon (0,\infty)^2 \to (0,\infty)$ by $g(x,y) \coloneqq x/(x+y)$, which is differentiable at $(p,1-p)$. Note that the distribution of (V_n,W_n) is an $N(0,\left(\begin{smallmatrix} p & 0 \\ 0 & q\end{smallmatrix}\right))$ distribution. By the delta method we find

$$\begin{aligned}
\sqrt{n}\big(U_{\lceil np\rceil} - p\big) &\stackrel{\mathrm{d}}{=} \sqrt{n}\left(g\left(\frac{V_n}{n}, \frac{W_n}{n}\right) - g(p,q)\right) \\
&\xrightarrow{\mathrm{d}} g'(p,1-p)N\left(0, \begin{pmatrix} p & 0 \\ 0 & q\end{pmatrix}\right) \\
&= N\left(0, g'(p,1-p)\begin{pmatrix} p & 0 \\ 0 & q\end{pmatrix} g'(p,1-p)^\top\right) \\
&= N(0,p(1-p)).
\end{aligned}$$

□

We now relate what we know about the uniform distribution to the quantile function:

Theorem 1.19. *Let $p \in (0,1)$ and let $X_1,\dots,X_n \stackrel{\text{iid}}{\sim} F$. Suppose that F is differentiable at $\xi_p \coloneqq F^{-1}(p)$ with derivative $f(\xi_p)$. Then*

$$\sqrt{n}\big(X_{(\lceil np\rceil)} - \xi_p\big) \xrightarrow{\mathrm{d}} N\left(0, \frac{p(1-p)}{f(\xi_p)^2}\right).$$

Proof. Let $U_1,\dots,U_n \stackrel{\text{iid}}{\sim} U(0,1)$, then we know that $F^{-1}(U_i) \stackrel{\mathrm{d}}{=} X_i$ and thus $F^{-1}(U_{(\lceil np\rceil)}) \stackrel{\mathrm{d}}{=} X_{(\lceil np\rceil)}$. Applying the delta method with $g = F^{-1}$, together with the previous theorem yields

$$\sqrt{n}\big(X_{(\lceil np\rceil)} - \xi_p\big) = \sqrt{n}\big(F^{-1}(U_{(\lceil np\rceil)}) - F^{-1}(p)\big) \xrightarrow{\mathrm{d}} (F^{-1})'(p)\cdot N(0,p(1-p)).$$

Noting that $(F^{-1})'(p) = \frac{1}{f(\xi_p)}$ yields the result. □

1.4 Concentration inequalities

We turn our attention to concentration inequalities, with a focus on finite-sample results (instead of results that only hold for $n \to \infty$).

Definition 1.20. A random variable X with mean 0 is called *sub-Gaussian* with parameter σ^2 if

$$M_X(t) = \mathbb{E}(e^{tX}) \le e^{t^2\sigma^2/2}$$

for every $t \in \mathbb{R}$.

Note that equality holds when $X \sim N(0,\sigma^2)$, since the MGF of an $N(\mu,\sigma^2)$ distribution is given by $t \mapsto \exp\big(\mu t + \sigma^2 t^2/2\big)$.

Recap 1.21. Recall the *tail bound formula* for the expectation: if X is a nonnegative random variable, then

$$\mathbb{E}[X] = \int_0^\infty \mathbb{P}(X > x)\,\mathrm{d}x\,.$$

Furthermore, recall that the *gamma function* is defined for $z \in (0, \infty)$ by

$$\Gamma(z) = \int_0^\infty x^{z-1} e^{-x} \, \mathrm{d}x$$

and satisfies $\Gamma(n) = (n-1)!$ for all $n \in \mathbb{N}$.

Finally, recall the following inequality: for all $a, b \in \mathbb{R}$ and $p \geq 1$

$$(a+b)^p \leq 2^{p-1}(a^p + b^p).$$

This follows from the convexity of the function $x \mapsto x^p$.

Proposition 1.22. *We consider some characterisations of sub-Gaussianity:*

(a) Let X be sub-Gaussian with parameter σ^2. Then

$$\max\{\mathbb{P}(X \geq x), \mathbb{P}(X \leq -x)\} \leq e^{-x^2/(2\sigma^2)} \quad \text{for every } x \geq 0. \tag{1}$$

(b) Let X be a random variable which satisfies $\mathbb{E}(X) = 0$ and eq. (1). Then for every $q \in \mathbb{N}$ it holds that

$$\mathbb{E}(X^{2q}) \leq 2 \cdot q!(2\sigma^2)^q \leq q!(2\sigma)^{2q}.$$

(c) If X is a random variable with $\mathbb{E}(X) = 0$ and $\mathbb{E}(X^{2q}) \leq q!C^{2q}$ for all $q \in \mathbb{N}$, then X is sub-Gaussian with parameter $4C^2$.

Proof. (a) We first consider $\mathbb{P}(X \geq x)$. By Markov's inequality, we have for all $t \in \mathbb{R}$ that

$$\mathbb{P}(X \geq x) = \mathbb{P}(e^{tX} \geq e^{tx}) \leq e^{-tX}\mathbb{E}(e^{tX}) \leq e^{-tx+t^2\sigma^2/2}.$$

Since the LHS is independent of t, we can take the infimum over t on the RHS and obtain

$$\mathbb{P}(X \geq x) \leq \inf_{t \in \mathbb{R}} e^{-tx+t^2\sigma^2/2} = e^{-x^2/(2\sigma^2)},$$

since the infimum of $t^2\sigma^2/2 - tx$ is attained at $t = x/\sigma^2$ (this method is called *Chernoff bounding*).

For $\mathbb{P}(X \leq -x) = \mathbb{P}(-X \geq x)$ we can use the fact that $-X$ is also sub-Gaussian with parameter σ^2.

(b) By the previous part, we have $\mathbb{P}(|X| \geq x) \leq 2e^{-x^2/(2\sigma^2)}$. Some calculations give

$$\begin{aligned}
\mathbb{E}(X^{2q}) &= \int_0^\infty \mathbb{P}(X^{2q} \geq x) \, \mathrm{d}x = \int_0^\infty \mathbb{P}(|X| \geq x^{1/(2q)}) \\
&= 2q \int_0^\infty x^{2q-1} \mathbb{P}(|X| \geq x) \, \mathrm{d}x \\
&\leq 4q \int_0^\infty x^{2q-1} e^{-x^2/(2\sigma^2)} \, \mathrm{d}x \, .
\end{aligned}$$

Now set $t = x^2/2\sigma^2$, so that $x = \sigma(2t)^{1/2}$ and thus $\mathrm{d}x = \sigma(2t)^{-1/2} \, \mathrm{d}t$. Plugging that in we get

$$\begin{aligned}
\mathbb{E}(X^{2q}) \leq 4q \int_0^\infty (\sigma(2t)^{1/2})^{2q-1} e^{-t} \sigma(2t)^{-1/2} \, \mathrm{d}t &= 2^{q+1} q \sigma^{2q} \int_0^\infty t^{q-1} e^{-t} \, \mathrm{d}t \\
= 2^{q+1} q \sigma^{2q} \Gamma(q) = 2 \cdot q!(2\sigma)^q.
\end{aligned}$$

(c) Note that $x \mapsto e^{-tx}$ is convex for every $t \in \mathbb{R}$, so $\mathbb{E}(e^{-tX}) \geq e^{-t\mathbb{E}(X)} = e^0 = 1$ by Jensen's inequality. Let X' denote an independent copy of X: then $X - X'$ has a symmetric distribution, so all its odd moments vanish. Therefore we find

$$\begin{aligned}
\mathbb{E}[e^{tX}] \leq \mathbb{E}[e^{-tX'}]\mathbb{E}[e^{tX}] &= \mathbb{E}[e^{t(X-X')}] = \mathbb{E}\sum_{q=0}^{\infty}\left[\frac{t^{2q}(X-X')^{2q}}{(2q)!}\right] \\
&= \sum_{q=0}^{\infty}\frac{t^{2q}\mathbb{E}[(X-X')^{2q}]}{(2q)!} \leq \sum_{q=0}^{\infty}\frac{2^{2q-1}t^{2q}\left(\mathbb{E}[X^{2q}]+\mathbb{E}[(X')^{2q}]\right)}{(2q)!} \\
&\leq \sum_{q=0}^{\infty}\frac{2^{2q-1}t^{2q}2q!C^{2q}}{(2q)!} = \sum_{q=0}^{\infty}\frac{(2tC)^{2q}q!}{(2q)!} = \sum_{q=0}^{\infty}\frac{(2tC)^{2q}}{\prod_{j=1}^{q}(q+j)} \\
&\leq \sum_{q=0}^{\infty}\frac{(2tC)^{2q}}{\prod_{j=1}^{q}(2j)} = \sum_{q=1}^{\infty}\frac{(2t^2C^2)^q}{q!} = e^{2t^2C^2}.
\end{aligned}$$

This shows that X is sub-Gaussian with parameter $4C^2$.

□

Note that the proposition is not an "if and only if"-type theorem: suppose we start with a sub-Gaussian variable X with parameter σ^2. Then by (b), we have $\mathbb{E}[X^{2q}] \leq q!(2\sigma)^{2q}$, and (c) then implies that X is sub-Gaussian with parameter $16\sigma^2$.

Theorem 1.23 (Hoeffding's inequality)**.** *Let $X_1, \ldots, X_n$ be independent sub-Gaussian random variables, with X_i having parameter σ_i^2. Then $\bar{X}$ is sub-Gaussian with parameter $\overline{\sigma^2}$. In particular, we have*

$$\max\left\{\mathbb{P}(\bar{X} \geq x), \mathbb{P}(\bar{X} \leq -x)\right\} \leq e^{-nx^2/(2\overline{\sigma^2})}.$$

Proof. For $t \in \mathbb{R}$, we have

$$\mathbb{E}[e^{t\bar{X}}] = \mathbb{E}[e^{(t/n)\sum_i X_i}] = \prod_{i=1}^{n}\mathbb{E}[e^{(t/n)X_i}] \leq \prod_{i=1}^{n}e^{t^2\sigma_i^2/(2n^2)} = e^{t^2\overline{\sigma^2}/(2n)},$$

which shows $\bar{X}$ is sub-Gaussian with parameter $\overline{\sigma^2}/n$. Applying part (a) of the previous proposition shows the second result. □

Remark. A direct consequence of Hoeffding's inequality is that

$$\mathbb{P}(\left|\bar{X}\right| \geq x) \leq 2e^{-nx^2/(2\overline{\sigma^2})}.$$

The inequality is often stated in this weaker way.

Lemma 1.24 (Hoeffding's lemma)**.** *Let X be a random variable with $\mathbb{E}X = 0$ that satisfies $a \leq X \leq b$. Then X is sub-Gaussian with parameter $(b-a)^2/4$.*

Proof. See Example Sheet 1, question 2. □

Corollary 1.25. *Let $X_1, \ldots, X_n$ be independent random variables where $\mathbb{E}[X_i] = \mu_i$ and $a_i \leq X_i \leq b_i$. Then we have*

$$\mathbb{P}(\bar{X} - \bar{\mu} \geq x) \leq \exp\left(-\frac{2n^2x^2}{\sum_{i=1}^{n}(b_i-a_i)^2}\right).$$

Proof. By Hoeffding's lemma, $X_i - \mu_i$ is sub-Gaussian with parameter $(b_i - a_i)^2/4$ for each i. The result now follows from theorem 1.23. □

Note that when X takes values in $[a, b]$, its variance is at most $(b-a)^2$. However, when $\mathrm{Var}(X_i) \ll (b_i - a_i)^2$, Hoeffding's inequality can be loose (for example, when $X_i \sim \mathrm{Bern}(p_i)$ with p_i small). In such circumstances, Bennett's or Bernstein's inequality may give better results.

Theorem 1.26 (Bennett's inequality)**.** *Let $X_1, \dots, X_n$ be independent random variables with $\mathbb{E}[X_i] = 0$, $\sigma_i^2 := \mathrm{Var}(X_i) < \infty$, and $X_i \le b$ for some $b > 0$. Define $S := \sum_{i=1}^n X_i$, $\nu := \overline{\mathrm{Var}(X_i)}$ and $\varphi\colon \mathbb{R} \to \mathbb{R}$ by $\varphi(u) := e^u - 1 - u = \sum_{k=2}^\infty \frac{u^k}{k!}$, then for every $t > 0$ we have*

$$\log \mathbb{E}[e^{tS}] \le \frac{n\nu}{b^2}\varphi(bt).$$

Defining $h\colon (0, \infty) \to [0, \infty)$ by $h(u) := (1+u)\log(1+u) - u$, we have for every $x > 0$ that

$$\mathbb{P}(\bar{X} \ge x) \le \exp\left(-\frac{n\nu}{b^2} h\left(\frac{bx}{\nu}\right)\right).$$

Proof. Define $g\colon \mathbb{R} \to \mathbb{R}$ by

$$g(u) := \sum_{k=0}^\infty \frac{u^k}{(k+2)!} = \begin{cases} \frac{\varphi(u)}{u^2} & \text{if } u \ne 0, \\ \frac{1}{2} & \text{if } u = 0. \end{cases}$$

Then one can check that g is increasing on $\mathbb{R}$, so

$$e^{tX_i} - 1 - tX_i = t^2 X_i^2 g(tX_i) \le t^2 X_i^2 g(tb) = X_i^2 \frac{\varphi(bt)}{b^2},$$

and therefore

$$e^{tX_i} \le 1 + tX_i + X_i^2 \frac{\varphi(bt)}{b^2} \implies \mathbb{E}[e^{tX_i}] \le 1 + \mathbb{E}[X_i^2]\frac{\varphi(bt)}{t^2} = 1 + \mathrm{Var}(X_i)\frac{\varphi(bt)}{b^2}.$$

Hence for $t > 0$ we have

$$\begin{aligned} \log \mathbb{E}[e^{tS}] &= \sum_{i=1}^n \log \mathbb{E}[e^{tX_i}] \le n \cdot \frac{1}{n}\sum_{i=1}^n \log\left(1 + \mathrm{Var}(X_i)\frac{\varphi(bt)}{b^2}\right) \\ &\overset{*}{\le} n \log\left(1 + \frac{\nu\varphi(bt)}{b^2}\right) \overset{**}{\le} \frac{n\nu}{b^2}\varphi(bt). \end{aligned}$$

Here, $(*)$ follows from the fact that log is a concave function while $(**)$ follows from the fact that $\log(1+u) \le u$ for all $u \ge 0$. This concludes the proof for the first part of the theorem.

Now, we apply the method of Chernoff bounding and find

$$\mathbb{P}(\bar{X} \ge x) = \mathbb{P}(S \ge nx) \le \inf_{t>0} e^{-ntx}\mathbb{E}[e^{tS}] \le \inf_{t>0} e^{-ntx + n\nu\varphi(bt)/b^2} = \exp\left(-\frac{n\nu}{b^2} h\left(\frac{bx}{\nu}\right)\right),$$

since once can check that the infimum is attained at $t = b^{-1}\log(1 + bx/\nu)$. □

Definition 1.27. A random variable X with $\mathbb{E}X = 0$ is called *sub-Gamma in the right tail* with variance factor $\sigma^2 > 0$ and scale $c > 0$ if

$$\mathbb{E}[e^{tX}] \le \exp\left(\frac{\sigma^2 t^2}{2(1 - ct)}\right)$$

for all $t \in [0, 1/c)$.

Note that this definition looks like that of sub-Gaussianity, except that e^{tX} can explode as t approaches $1/c$. We give some characteristics of sub-Gamma distributions:

Definition 1.28. For any $x \in \mathbb{R}$ we define $x_+ \coloneqq \max(x, 0)$.

Proposition 1.29. *(a) Let X be sub-Gamma in the right tail with variance factor σ^2 and scale c. Then*

$$\mathbb{P}(X \geq x) \leq \exp\left(-\frac{x^2}{2(\sigma^2 + cx)}\right)$$

for all $x \geq 0$.

(b) Let X be a random variable with $\mathbb{E}X = 0$, $\mathbb{E}[X^2] \leq \sigma^2$ and $\mathbb{E}[(X_+)^q] \leq q!\sigma^2 c^{q-2}/2$ for all $q \geq 3$. Then X is sub-Gamma in the right tail with variance factor σ^2 and scale parameter c.

Proof. (a) Again, we apply a Chernoff bound: we have

$$\mathbb{P}(X \geq x) \leq \inf_{t \in [0,1/c)} e^{-tx}\mathbb{E}[e^{tX}] \leq \inf_{t \in [0,1/c)} \exp\left(-tx + \frac{\sigma^2 t^2}{2(1-ct)}\right)$$
$$\leq \exp\left(-\frac{x^2}{2(\sigma^2 + cx)}\right),$$

where we have set $t = x/(\sigma^2 + cx) \in [0, 1/c)$ in the final step.

(b) Recall from the proof of Bennett's inequality that g is increasing and therefore for $u \leq 0$ we have $\varphi(u) = u^2 g(u) \leq u^2 g(0) = \frac{u^2}{2}$. Therefore, for every $u \in \mathbb{R}$ we have

$$\varphi(u) \leq \frac{u^2}{2} + \sum_{q=3}^{\infty} \frac{(u_+)^q}{q!}.$$

We deduce that for $t \in [0, 1/c)$ we have (note $\log(x) \leq x - 1$ for all x):

$$\log \mathbb{E}[e^{tX}] \leq \mathbb{E}(e^{tX}) - 1 = \mathbb{E}[\varphi(tX)] \leq \mathbb{E}\left[\frac{t^2 X^2}{2} + \sum_{q=3}^{\infty} \frac{t^q X_+^q}{q!}\right].$$

By Fubini's theorem, since the infinite sum has only positive terms we may interchange sum and expectation to obtain

$$\mathbb{E}\left[\frac{t^2X^2}{2} + \sum_{q=3}^{\infty}\frac{t^q \mathbb{E}[X_+^q]}{q!}\right] = \frac{t^2 \operatorname{Var}[X]}{2} + \sum_{q=3}^{\infty}\frac{t^q\mathbb{E}[X_+^q]}{q!} \leq \frac{\sigma^2 t^2}{2}\sum_{q=2}^{\infty} t^{q-2}c^{q-2} = \frac{\sigma^2 t^2}{2(1-ct)}.$$

□

Following this proposition, we can prove Bernstein's inequality:

Theorem 1.30 (Bernstein's inequality)**.** *Let $X_1, \ldots, X_n$ be independent random variables with $\mathbb{E}[X] = 0$, $\frac{1}{n}\sum_{i=1}^n \operatorname{Var}(X_i) \leq \sigma^2$ and $\frac{1}{n}\sum_{i=1}^n \mathbb{E}[(X_i)_+^q] \leq q!\sigma^2 c^{q-2}/2$ some $\sigma, c > 0$ and for all $q \geq 3$. Then $S \coloneqq \sum_{i=1}^n X_i$ is sub-Gamma in the right tail with variance factor $n\sigma^2$ and scale parameter c. In particular we have*

$$\mathbb{P}(\bar{X} \geq x) \leq \exp\left(-\frac{nx^2}{2(\sigma^2 + cx)}\right),$$

for all $x \geq 0$.

Proof. We have by part (b) of the previous proposition

$$\log \mathbb{E}[e^{tS}] = \sum_{i=1}^{n} \log \mathbb{E}[e^{tX_i}] \leq n\frac{\sigma^2 t^2}{2(1-ct)},$$

and the second claim follows from part (a) of the previous proposition:

$$\mathbb{P}(\bar{X} \geq x) = \mathbb{P}(S \geq nx) \leq \exp\left(-\frac{nx^2}{2(\sigma^2 + cx)}\right).$$

$\square$

2 Kernel density estimation

Let $X_1, \dots, X_n \overset{\text{iid}}{\sim} f$, and suppose we wish to estimate the density function f. The oldest way to do this is with a histogram: we divide $\mathbb{R}$ into equally sized intervals or *bins*, and let I_x denote the bin containing $x \in \mathbb{R}$.

Definition 2.1. The *histogram density estimator* $\hat{f}_n^H$ with bin width $b > 0$ is given by

$$\hat{f}_n^H(x) \coloneqq \frac{1}{nb} \sum_{i=1}^{n} \mathbb{1}_{X_i \in I_x}.$$

There are a few major drawbacks to using histograms: it is difficult to choose b and the positioning of bin edges, the theoretical performance is suboptimal (mostly due to their discontinuity) and graphical display in the multivariate case is difficult.

2.1 The univariate kernel density estimator

Definition 2.2. A Borel measurable function $K\colon \mathbb{R} \to \mathbb{R}$ is called a *kernel* if it satisfies $\int_{\mathbb{R}} K(x)\,\mathrm{d}x = 1$.

A *univariate kernel density estimator* of f with kernel K and *bandwidth* $h > 0$ is defined as

$$\hat{f}_n(x) \coloneqq \frac{1}{nh} \sum_{i=1}^{n} K\left(\frac{x - X_i}{h}\right).$$

Defining $K_h(x) \coloneqq \frac{1}{h} K\left(\frac{x}{h}\right)$, we can rewrite this as

$$\hat{f}_n(x) \coloneqq \frac{1}{n} \sum_{i=1}^{n} K_h(x - X_i).$$

Usually K is chosen to be non-negative (which ensures that K itself and $\hat{f}_n$ are themselves density functions), and K is often chosen to be symmetric about 0. Generally, the choice of kernel K is much less important than the choice of bandwidth h.

If we consider $\hat{f}_n(x)$ as a point estimator of $f(x)$, we typically wish to minimise the *mean squared error*

$$\operatorname{MSE}(\hat{f}_n(x)) \coloneqq \mathbb{E}\Big[(\hat{f}_n(x) - f(x))^2\Big].$$

Other possibilities include the mean absolute error which (unlike the MSE) is scale-invariant. However, the MSE has an appealing decomposition into variance and bias terms:

$$\operatorname{MSE}(\hat{f}_n(x)) = \operatorname{Var}(\hat{f}_n(x)) + \operatorname{Bias}^2(\hat{f}_n(x)).$$

We can express the MSE in terms of convolutions:

Definition 2.3. Let $g_1, g_2\colon \mathbb{R} \to \mathbb{R}$ be measurable. Then the *convolution* of g_1 and g_2, denoted $g_1 * g_2$, is defined by

$$(g_1 * g_2)(x) \coloneqq \int_{\mathbb{R}} g_1(x - z) g(z)\,\mathrm{d}z\,.$$

We can compute

$$\begin{aligned}\operatorname{Bias} \hat{f}_n(x) &= \mathbb{E}[\hat{f}_n(x)] - f(x) = \mathbb{E}[K_h(x - X_1)] - f(x) = \int_{\mathbb{R}} K_h(x - z) f(z)\,\mathrm{d}z \\ &= (K_h * f)(x) - f(x). \end{aligned} \tag{2}$$

Analogously, letting $\xi_i \coloneqq K_h(x - X_i)$ (note that these are i.i.d. random variables), we have

$$\operatorname{Var} \hat{f}_n(x) = \frac{1}{n} \operatorname{Var}(\xi_1) = \frac{1}{n}\big(\mathbb{E}[\xi_1^2] - \mathbb{E}^2[\xi_1]\big) = \frac{1}{n}\big[(K_h^2 * f)(x) - (K_h * f)^2(x)\big]. \tag{3}$$

To assess performance of h and K, we want to assess the performance of $\hat{f}_n$ as an estimation of f as a function. This gives the following definition:

Definition 2.4. We define the *mean integrated squared error* or MISE as

$$\operatorname{MISE}(\hat{f}_n) \coloneqq \mathbb{E}\left(\int_{\mathbb{R}} \Big(\hat{f}_n(x) - f(x)\Big)^2 \mathrm{d}x\right) \overset{\star}{=} \int_{\mathbb{R}} \operatorname{MSE}(\hat{f}_n(x)) \,\mathrm{d}x\,,$$

where $\star$ follows from Fubini's theorem since the integrand is nonnegative.

We now aim to find bounds on the bias and the variance of $\hat{f}_n$ in order to choose h and K appropriately.

2.2 Bounds on variance and bias

Definition 2.5. For a kernel K, define $R(K) \coloneqq \int_{\mathbb{R}} K^2(u)\,\mathrm{d}u$.

Proposition 2.6. *Let $\hat{f}_n$ be the kernel density estimator with kernel K and bandwidth $h > 0$ constructed from $X_1, \dots, X_n \overset{\text{iid}}{\sim} f$. Then for any $x \in \mathbb{R}, h > 0, n \in \mathbb{N}$ we have*

$$\operatorname{Var} \hat{f}_n(x) \leq \frac{1}{nh} \|f\|_\infty R(K).$$

Proof. By eq. (3) we have

$$\begin{aligned} \operatorname{Var} \hat{f}_n(x) &\leq \frac{1}{n}(K_h^2 * f)(x) = \frac{1}{nh^2}\int_{\mathbb{R}} K^2\left(\frac{x-z}{h}\right) f(z)\,\mathrm{d}z = \frac{1}{nh}\int_{\mathbb{R}} K^2(u) f(x - uh)\,\mathrm{d}u \qquad (4) \\ &\leq \frac{1}{nh}\|f\|_\infty \int_{\mathbb{R}} K^2(u)\,\mathrm{d}u = \frac{1}{nh}\|f\|_\infty R(K). \end{aligned}$$

□

Obtaining a bound on the bias is not at all straightforward: we wil need to introduce conditions on both the density f and the kernel K.

Definition 2.7. Let $I \subseteq \mathbb{R}$ be an interval, fix $\beta, L > 0$, and let $m \coloneqq \lceil \beta \rceil - 1$. A function $f \colon \mathbb{R} \to \mathbb{R}$ is said to belong to the *Hölder class* $\mathcal{H}(\beta, L)$ if f is m times differentiable on I and

$$\left| f^{(m)}(x) - f^{(m)}(y) \right| \leq L|x - y|^{\beta - m} \quad \text{for all } x, y \in I.$$

If I is unspecified, we let $I = \mathbb{R}$.

The densities in $\mathcal{H}(\beta, L)$ are denoted by

$$\mathcal{F}(\beta, L) \coloneqq \left\{ f \in \mathcal{H}(\beta, L) \mid f \geq 0 \text{ and } \int_{\mathbb{R}} f \,\mathrm{d}x = 1 \right\}.$$

Definition 2.8. Fix $\ell \in \mathbb{N}$. We say a kernel K is of *order* ℓ if $\int_{\mathbb{R}} x^j k(x)\,\mathrm{d}x = 0$ for $j = 1, \dots, \ell - 1$.

Remark. Most kernels used in practice are of order 2. Note that a kernel of order ≥ 3 cannot be nonnegative, since we have $\int_{\mathbb{R}} x^2 K(x)\,\mathrm{d}x = 0$. Therefore, the kernels are not themselves densities and the corresponding kernel density estimate is not guaranteed to be a density.

Proposition 2.9. *Assume that $f \in \mathcal{F}(\beta, L)$ and that K is a kernel of order $\ell := \lceil \beta \rceil$, and furthermore assume that*

$$\mu_\beta(K) := \int_{\mathbb{R}} |u|^\beta |K(u)| \,\mathrm{d}u < \infty.$$

Then the kernel density estimate with bandwidth h and kernel K based on $X_1, \dots, X_n \sim f$ satisfies

$$\left|\mathrm{Bias}\, \hat{f}_n(x)\right| \le \frac{L}{(\ell-1)!}\mu_\beta(K) h^\beta \quad \textit{for all } x \in \mathbb{R}, h > 0, n \in \mathbb{N}.$$

Proof. By eq. (2), we have

$$\mathrm{Bias}\, \hat{f}_n(x) = \frac{1}{h}\int_{\mathbb{R}} K\left(\frac{x-z}{h}\right) f(z)\,\mathrm{d}z - f(x) = \int_{\mathbb{R}} K(u)(f(x-uh) - f(x))\,\mathrm{d}x\,.$$

By applying Taylor's theorem with the Lagrange remainder we obtain, with $m = \lceil \beta \rceil - 1$, that

$$f(x-uh) - f(x) = \sum_{j=1}^{m-1} \frac{(-uh)^j}{j!} f^{(j)}(x) + \frac{(-uh)^m}{m!} f^{(m)}(x - \tau uh) \quad \text{for some } \tau \in [0,1].$$

Since $\int_{\mathbb{R}} u^j K(u)\,\mathrm{d}u = 0$ for all $j \le m$, plugging the sum into the integral will give 0. Therefore, we find

$$\mathrm{Bias}\, \hat{f}_n(x) = \frac{(-h)^m}{m!}\int_{\mathbb{R}} u^m K(u) f^{(m)}(x - \tau uh)\,\mathrm{d}u = \frac{(-h)^m}{m!}\int_{\mathbb{R}} u^m K(u)\left[f^{(m)}(x-\tau uh) - f^{(m)}(x)\right]\mathrm{d}u\,,$$

where the last inequality follows again from the fact that K is of order $m+1$.

Now we use that $f \in \mathcal{F}(\beta, L)$, and conclude

$$\left|\mathrm{Bias}\, \hat{f}_n(x)\right| \le \frac{Lh^m}{m!}\int_{\mathbb{R}} |u|^m |K(u)| |\tau uh|^{\beta - m}\,\mathrm{d}u \le \frac{Lh^\beta}{m!}\int_{\mathbb{R}} |u|^\beta |K(u)|\,\mathrm{d}u = \frac{L}{(\ell-1)!}\mu_\beta(K)h^\beta,$$

which concludes the proof. ◻

Combining propositions 2.6 and 2.9, we find that

$$\mathrm{MSE}\, \hat{f}_n(x) \le \frac{1}{nh} R(K)\|f\|_\infty + \frac{L^2}{((\ell-1)!)^2}\mu_\beta^2(K) h^{2\beta}.$$

By minimising this w.r.t. h, we find that the optimal h is given by

$$h_n^* = \left(\frac{((\ell-1)!)^2 \|f\|_\infty R(K)}{2\beta L^2 \mu_\beta^2(K)}\right)^{1/(2\beta+1)} n^{-1/(2\beta+1)},$$

and the corresponding MSE is given by

$$\mathrm{MSE}\, \hat{f}_n(x) \le \left(\frac{\|f\|_\infty^{2\beta} R(K)^{2\beta} L^2 \mu_\beta^2(K)\left[(2\beta)^{2\beta+1} + 1\right]}{((\ell-1)!)^2 (2\beta)^{2\beta}}\right) n^{-2\beta/(2\beta+1)},$$

This $O(n^{-2\beta/(2\beta+1)})$ bound on the rate is slower than the $O(1/n)$ rate found in parametric problems, but such a rate is only obtained when the assumed model is correct.

We can strengthen this as follows:

Theorem 2.10. *Assume that K is a kernel of order $\ell := \lceil \beta \rceil$ and that $\mu_\beta(K)$ and $R(K)$ are both finite. Fix $\alpha > 0$ and let $h = \alpha n^{-1/(2\beta+1)}$. Then there exists $C > 0$, independent of n, such that*

$$\sup_{x \in \mathbb{R}} \sup_{f \in \mathcal{F}(\beta, L)} \mathrm{MSE}\, \hat{f}_n(x) \le C n^{-2\beta/(2\beta+1)}.$$

Proof. We will show that the class $\mathcal{F}(\beta, L)$ is uniformly bounded in supremum norm. Let K^* be a bounded kernel of order ℓ (see example sheet TODO:), then by the previous proposition with $h = 1$ we have by nonnegativity of f that

$$\begin{aligned} f(x) &\leq \left| f(x) - \int_{-\infty}^{\infty} K^*(x-z) f(z)\,\mathrm{d}z \right| + \left| \int_{-\infty}^{\infty} K^*(x-z) f(z)\,\mathrm{d}z \right|. \\ &\leq \left| \operatorname{Bias} \hat{f}_{n,K^*}(x) \right| + \|K^*\|_\infty \int_{-\infty}^{\infty} f(z)\,\mathrm{d}z \\ &\leq \frac{L}{(\ell-1)!} \mu_\beta(K^*) + \|K^*\|_\infty, \end{aligned}$$

and this bound is independent of f and x.

Now we have

$$\operatorname{MSE} \hat{f}_n(x) \leq \frac{R(K)\|f\|_\infty}{nh} + \frac{L^2}{((\ell-1)!)^2} \mu_\beta^2(K) h^{2\beta} \leq C n^{-2\beta/(2\beta+1)}.$$

□

2.3 Bounds on the integrated variance and bias

To bound the MISE, we will give bounds on the integrated variance and bias.

Proposition 2.11. *Let $\hat{f}_n$ denote the kernel density estimate with bandwidth h and kernel K based on $X_1, \ldots, X_n \overset{\text{iid}}{\sim} P$ (where P is a distribution on $\mathbb{R}$). Then*

$$\int_{-\infty}^{\infty} \operatorname{Var} \hat{f}_n(x)\,\mathrm{d}x = \frac{1}{nh} R(K).$$

Proof. We have by Fubini and eq. (4) that

$$\begin{aligned} \int_{-\infty}^{\infty} \operatorname{Var} \hat{f}_n(x)\,\mathrm{d}x &\leq \frac{1}{nh^2} \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} K^2\left(\frac{x-z}{h}\right) f(z)\,\mathrm{d}z\,\mathrm{d}x = \frac{1}{nh^2} \int_{-\infty}^{\infty} f(z) \int_{-\infty}^{\infty} K^2\left(\frac{x-z}{h}\right) \mathrm{d}x\,\mathrm{d}z \\ &= \frac{1}{nh} R(K) \int_{-\infty}^{\infty} f(z)\,\mathrm{d}z = \frac{1}{nh} R(K). \end{aligned}$$

□

Recap 2.12. Let $[a, b] = I \subseteq \mathbb{R}$ be an interval, then $f \colon I \to \mathbb{R}$ is called *absolutely continuous* if, for every $\varepsilon > 0$, there exists $\delta > 0$ such that, whenever $(x_1, y_1), \ldots, (x_m, y_m)$ are disjoint subintervals of I with $\sum_{i=1}^m (y_i - x_i) < \delta$, we have $\sum_{i=1}^m |f(y_i) - f(x_i)| < \varepsilon$.

It is known that absolute continuity is equivalent to being differentiable Lebesgue almost everywhere with a so-called *weak derivative* f' that satisfies $f(x) = f(a) + \int_a^x f'(t)\,\mathrm{d}t$ for all $x \in [a, b]$.

Recap 2.13. The *generalised Minkowski inequality* states that any Borel measurable function $g \colon \mathbb{R}^2 \to \mathbb{R}$ we have that

$$\int_{\mathbb{R}} \left(\int_{\mathbb{R}} g(u, x)\,\mathrm{d}u \right)^2 \mathrm{d}x \leq \left(\int_{\mathbb{R}} \left(\int_{\mathbb{R}} g^2(u, x)\,\mathrm{d}x \right)^{1/2} \mathrm{d}u \right)^2.$$

To obtain bounds on the integrated squared bias, we will require smoothness conditions w.r.t. the $L^2(\mathbb{R})$ norm.

Definition 2.14. Fix $\beta, L > 0$ and let $m := \lceil \beta \rceil - 1$. The *Nikolski class* $\mathcal{N}(\beta, L)$ consists of functions $f\colon \mathbb{R} \to \mathbb{R}$ that are $(m-1)$ times differentiable and for which $f^{(m-1)}$ is absolutely continuous with weak derivative $f^{(m)}$ satisfying

$$\left\{ \int_{-\infty}^{\infty} \left[f^{(m)}(x+t) - f^{(m)}(x) \right]^2 \mathrm{d}x \right\}^{1/2} \le L|t|^{\beta - m} \quad \text{for all } t \in \mathbb{R}.$$

The densities in $\mathcal{N}(\beta, L)$ are denoted by $\mathcal{F}_{\mathcal{N}}(\beta, L)$.

Proposition 2.15. *Fix $\beta, L > 0$ and let K be a kernel of order $\ell := \lceil \beta \rceil$. Let $\hat{f}_n$ denote the KDE with kernel K and bandwidth h based on $X_1, \ldots, X_n \overset{\text{iid}}{\sim} f \in \mathcal{F}_{\mathcal{N}}(\beta, L)$. Then we have*

$$\int_{-\infty}^{\infty} \operatorname{Bias}^2 \hat{f}_n(x) \,\mathrm{d}x \le \frac{L^2}{((\ell-1)!)^2} \mu_\beta^2(K) h^{2\beta}.$$

Proof. TODO: write this out (integration + taylor expansion + 2x minkowski). □

Putting everything together, we obtain the following:

Theorem 2.16. *Fix $\beta, L > 0$, and let K be a kernel of order $\ell = \lceil \beta \rceil$ with $R(k)$ and $\mu_\beta(K)$ finite. Let $\hat{f}_n$ be the KDE with kernel K and bandwidth h based on $X_1, \ldots, X_n \overset{\text{iid}}{\sim} f \in \mathcal{F}_{\mathcal{N}}(\beta, L)$. Then we have*

$$\operatorname{MISE} \hat{f}_n \le \frac{R(K)}{nh} + \frac{L^2}{((\ell-1)!)^2} \mu_\beta^2(K) h^{2\beta}.$$

In particular, fixing $\alpha > 0$ and taking $h = \alpha n^{-1/(2\beta+1)}$, there exists $C > 0$ independent of n such that

$$\sup_{f \in \mathcal{F}_{\mathcal{N}}(\beta, L)} \operatorname{MISE} \hat{f}_{n,h,K} \le C n^{-2\beta/(2\beta+1)}.$$

2.4 Bandwidth selection

The choice of bandwidth in the previous theorem is not practical since we have not specified α and β is typically unknown.

2.4.1 Least squares cross validation

One possible approach is *least squares cross validation*. For this, note that minimising the MISE is equivalent to minimising

$$\operatorname{MISE}(\hat{f}_n) - \int_{\mathbb{R}} f^2(x) \,\mathrm{d}x = \mathbb{E}\left[\int_{\mathbb{R}} \hat{f}_n^2(x) \,\mathrm{d}x \right] - 2\mathbb{E}\left[\int_{\mathbb{R}} \hat{f}_n(x) f(x) \,\mathrm{d}x \right],$$

and it can be checked that an unbiased estimator for the above is given by

$$\operatorname{LSCV}(h) := \int_{\mathbb{R}} \hat{f}_n^2(x) \,\mathrm{d}x - \frac{2}{n} \sum_{i=1}^{n} \hat{f}_{n,-i}(X_i),$$

where $\hat{f}_{n,-i}$ is the KDE based on all observations except X_i. We now choose h such that LSCV is minimised.

2.4.2 Lepski

TODO: write this subsection (TODO: understand this first)

2.5 Choice of kernel

To choose a kernel, we first fix the scale of the kernel by setting $\mu_2(K) = 1$. Now, by our bound on the MISE (theorem 2.16) it is reasonable to minimise $R(K)$, where for simplicity we assume that K is a nonnegative second-order kernel. The solution is the Epanechnikov kernel

$$K_E(x) \coloneqq \frac{3}{4\sqrt{5}}\left(1 - \frac{x^2}{5}\right)\mathbb{1}_{|x|\leq\sqrt{5}},$$

and the ratio $R(K_E)/R(K)$ is called the *efficiency* of a kernel K. We find that for different kernels, the efficiency is greater than 0.9, which shows that kernel shape does not affect efficiency greatly.

2.6 Multivariate density estimation

The general d-dimensional KDE is

$$\hat{f}_n(x) \coloneqq \frac{1}{n\sqrt{\det H}} \sum_{i=1}^{n} K(H^{-1/2}(x - X_i)),$$

where H is a symmetric positive-definite bandwidth bandwidth matrix (often chosen to be diagonal or a multiple of I). If $H = h^2 I$, it can be shown that, under an appropriate definition of a β smoothness class, we have an optimal bandwidth of order $n^{-1/(d+2\beta)}$, and with this choise, a MISE of order $n^{-2\beta/(d+2\beta)}$. This is called the "curse of dimensionality": the higher the dimension becomes, the harder nonparametric estimation gets.

3 Nonparametric regression

3.1 Fixed and random design

In *fixed design*, we assume we have data $x_1 \leq \cdots \leq x_n$ and the response variables satisfy

$$Y_i \coloneqq m(x_i) + \sqrt{v(x_i)}\varepsilon_i,$$

where the ε_i are independent, mean-zero random variables with $\mathrm{Var}(\varepsilon_i) = 1$. The function m is called the *regression function*, and the function v is the *variance function*. If v is constant, the model is called *homoscedastic*, else it is called *heteroscedastic*.

In *random design*, we assume we have i.i.d. data pairs (X_i, Y_i) with

$$Y_i = m(X_i) + \sqrt{v(X_i)}\varepsilon_i,$$

where the ε_i are again independent with $\mathbb{E}[\varepsilon_1|X_1] = 0$ and $\mathrm{Var}(\varepsilon_1|X_1) = 1$. The regression function is given by $m(x) = \mathbb{E}(Y_1|X_1 = x)$ and the variance function by $v(x) = \mathrm{Var}(Y_1|X_1 = x)$.

3.2 Local polynomial estimators

We will assume the fixed design setting.

Definition 3.1. Let K be a kernel, $h > 0$ a bandwidth and $p \in \mathbb{N}$. Then the *local polynomial estimator of degree p with bandwidth h and kernel K*, denoted $\hat{m}_n(\cdot; p) \equiv \hat{m}_n(\cdot; p, h, K)$, is constructed at $x \in \mathbb{R}$ by fitting a polynomial p to the data using weighted least squares, where the pair (x_i, Y_i) is assigned weight $K_h(x_i - x)$.

To write this in formulas, define $Q(u) \coloneqq (1, u, \frac{u^2}{2}, \dots, \frac{u^p}{p!}) \in \mathbb{R}^{p+1}$ and $Q_h(\cdot) = Q(\cdot/h)$, we then have

$$\hat{m}_n(x; p) = \hat{\beta}_0, \quad \text{where } \hat{\beta} \coloneqq \operatorname*{arg\,min}_{\beta \in \mathbb{R}^{p+1}} \sum_{i=1}^{n} \big(Y_i - \beta^\top Q_h(x_i - x)\big)^2 K_h(x_i - x).$$

In matrix-vector notation, writing

$$X \equiv X(x; p, h) \coloneqq \begin{pmatrix} Q_h(x_1 - x)^\top \\ \vdots \\ Q_h(x_n - x)^\top \end{pmatrix} \in \mathbb{R}^{n \times (p+1)}, \quad Y \coloneqq \begin{pmatrix} Y_1 \\ \vdots \\ Y_n \end{pmatrix} \in \mathbb{R}^n,$$
$$W \equiv W(x; h, K) \coloneqq \mathrm{diag}(K_h(x_1 - x), \dots, K_h(x_n - x)) \in \mathbb{R}^{n \times n},$$

we have

$$\hat{\beta} = \operatorname*{arg\,min}_{\beta \in \mathbb{R}^{p+1}} (Y - X\beta)^\top W (Y - X\beta).$$

By standard weighted least squares theory, we know that $\hat{\beta}$ must satisfy $X^\top W X \hat{\beta} = X^\top W Y$.

Proposition 3.2. *Suppose $X^\top W X$ is positive definite. Then we have*

$$\hat{\beta} = (X^\top W X)^{-1} X^\top W Y.$$

We will assume from here on that $X^\top W X$ is indeed positive definite. In this case, since the entries of W and X are functions of $x_i - x$, we can write the local polynomial estimator in the form

$$\hat{m}_n(x) = n^{-1} \sum_{i=1}^{n} w(x, x_i) Y_i.$$

(??) The set of weights $\{w(x, x_i)\}$ is called the *effective kernel* at x.

For $p = 0$ and $p = 1$, there exist explicit formulas for the local polynomial estimator of degree p.

Proposition 3.3 (Reproducing property)**.** *Let $\{w_p(x, x_i)\}$ denote the effective kernel of a local polynomial estimator of degree p based on data $(x_1, Y_1), \ldots, (x_n, Y_n)$, and let R denote a polynomial of degree at most p. If $X^\top WX$ is positive definite, then*

$$\frac{1}{n}\sum_{i=1}^{n} w_p(x, x_i)R(x_i) = R(x).$$

Proof. We have

$$\frac{1}{n}\sum_{i=1}^{n} w_p(x, x_i)R(x_i) = \hat{\beta}_0,$$

where

$$\hat{\beta} = \underset{\beta\in\mathbb{R}^{p+1}}{\arg\min} \sum_{i=1}^{n} \big(R(x_i) - \beta^\top Q_h(x_i - x)\big)^2 K_h(x_i - x).$$

By choosing $\hat{\beta}$ such that $\hat{\beta}^\top Q_h(\cdot - x) = R$, we obtain 0, and therefore this is indeed the minimiser. Since $Q_h(0) = e_1$, it follows that $R(x) = \hat{\beta}^\top Q_h(x - x) = \hat{\beta}_0$, which concludes the proof. □

Before we can study the bias and variance of local polynomial estimators, we require the following lemma:

Lemma 3.4. *Let K be a kernel that vanishes outside $[-1, 1]$, and suppose that $n^{-1}X^\top WX$ is positive definite with minimal eigenvalue $\lambda_0 \equiv \lambda_{0,n,x} > 0$. Then*

(i) $\sup_{x\in[0,1]} \max_{i=1,\ldots,n} \frac{1}{n}|w(x, x_i)| \le \frac{2\|K\|_\infty}{\lambda_0 nh}$*;*

(ii) $n^{-1}\sum_{i=1}^{n} |w(x, x_i)| \le \frac{2\|K\|_\infty}{\lambda_0 nh} \cdot \sum_{i=1}^{n} \mathbb{1}_{|x_i-x|\le h}$*;*

(iii) $w(x, x_i) = 0$ *when* $|x_i - x| > h$.

Proof. (i) Note that $n^{-1}w(x, x_i)$ is the $(0, i)$ entry of the matrix $(X^\top WX)^{-1}X^\top W$, and it is therefore less than the norm of the i-th column of $(X^\top WX)^{-1}X^\top W$. For $x \in [0, 1]$ and $i = 1, \ldots, n$, we find

$$\begin{aligned}\frac{1}{n}|w(x, x_i)| &\le \big\|(X^\top WX)^{-1}Q_h(x_i - x)K_h(x_i - x)\big\| \overset{\star}{\le} \|K_h\|_\infty \big\|(X^\top WX)\big\|^{-1} \|Q_h(x_i - x)\|\mathbb{1}_{|x_i-x|\le h} \\ &= \frac{\|K\|_\infty}{h}\frac{1}{\lambda_0 n}\|Q_h(x_i - x)\|\mathbb{1}_{|x_i-x|\le h} \le \frac{\|K\|_\infty}{\lambda_0 nh}\|Q(1)\| = \frac{\|K\|_\infty}{\lambda_0 nh}\left(\sum_{j=0}^{p}\frac{1}{(j!)^2}\right)^{1/2} \\ &\le \frac{\|K\|_\infty}{\lambda_0 nh}e^{1/2} \le \frac{2\|K\|_\infty}{\lambda_0 nh}.\end{aligned}$$

Here, the indicator function in $\star$ appears because K vanishes outside $[-1, 1]$. Since the upper bound is independent of both x and i, the claim is proven.

(ii) Similarly as before, we find

$$\frac{1}{n}\sum_{i=1}^{n} |w(x, x_i)| \le \frac{\|K\|_\infty}{\lambda_0 nh}\sum_{i=1}^{n} \|Q_h(x_i - x)\|\mathbb{1}_{|x_i-x|\le h} \le \frac{2\|K\|_\infty}{\lambda_0 nh}\sum_{i=1}^{n} \mathbb{1}_{|x_i-x|\le h}.$$

(iii) This follows immediately from inequality $\star$ in the proof of (i).

□

Now, we can compute bounds on the variance and bias of our local polynomial estimator. For simplicity, we will assume $x_i = i/n$.

Proposition 3.5. *Assume the model $Y_i = m(x_i) + v^{1/2}\varepsilon_i$ with $m \in \mathcal{H}(\beta, L)$ on $[0,1]$ and $\max_i v(x_i) \leq \sigma^2_{\max}$. Let K be a kernel that vanishes outside $[-1,1]$, and suppose that λ_0, the minimal eigenvalue of $n^{-1}X^\top WX$, is strictly positive. Then, for $p \geq \lceil\beta\rceil - 1 =: \beta_0$ and for each $x_0 \in [0,1]$, $n \in \mathbb{N}$ and $h \geq 1/(2n)$, we have*

$$\operatorname{Var}\hat{m}_n(x_0;p) \leq \frac{16\|K\|_\infty^2\sigma^2_{\max}}{\lambda_0^2 nh}, \quad |\operatorname{Bias}\hat{m}_n(x_0;p)| \leq \frac{8L\|K\|_\infty}{\lambda_0\beta_0!}h^\beta.$$

Proof. Using the previous lemma, we obtain

$$\begin{aligned}\operatorname{Var}\hat{m}_n(x_0;p) &= \operatorname{Var}\left(\frac{1}{n}\sum_{i=1}^n w(x_0,x_i)Y_i\right) = \frac{1}{n^2}\sum_{i=1}^n w(x_0,x_i)^2\operatorname{Var}(Y_i) \\ &\leq \frac{\sigma^2_{\max}}{n^2}\sum_{i=1}^n w(x_0,x_i)^2 \leq \sigma^2_{\max}\left(\sup_{x\in[0,1]}\max_i \frac{1}{n}|w(x,x_i)|\right)\frac{1}{n}\sum_{i=1}^n |w(x_0,x_i)| \\ &\leq \frac{4\|K\|_\infty^2\sigma^2_{\max}}{\lambda_0^2 n^2h^2}\sum_{i=1}^n \mathbb{1}_{|i/n-x_0|\leq h}.\end{aligned}$$

Now, we have $|i/n - x_0| \leq h \iff i \in [nx_0 - nh, nx_0 + nh]$, and there are at most $2nh+1$ integers in this interval. Recalling that $1 \leq 2nh$ we obtain

$$\operatorname{Var}\hat{m}_n(x_0;p) \leq \frac{4(2nh+1)\|K\|_\infty\sigma^2_{\max}}{\lambda_0^2n^2h^2} \leq \frac{16\|K\|_\infty\sigma^2_{\max}}{\lambda_0^2 nh}.$$

For the bias, we will first use a Taylor expansion combined with the reproducing property proposition 3.3. Firstly, since $\frac{1}{n}\sum_{i=1}^n w(x_0,x_i) = 1$ by this property, we have

$$\operatorname{Bias}\hat{m}_n(x_0;p) = \left(\frac{1}{n}\sum_{i=1}^n w(x_0,x_i)m(x_i)\right) - m(x_0) = \frac{1}{n}\sum_{i=1}^n w(x_0,x_i)\{m(x_i) - m(x_0)\}.$$

Now, we apply a Taylor expansion to write

$$m(x_i) - m(x_0) = P(x_i - x_0) + \frac{1}{\beta_0!}m^{(\beta_0)}(x_0 + \tau_i(x_i - x_0))(x_i - x_0)^{\beta_0},$$

where P has degree at most $\beta_0 - 1 < p$ and a constant coefficient equal to 0, and $\tau_i \in [0,1]$. By the reproducing property we have $n^{-1}\sum_{i=1}^n w(x_0,x_i)P(x_i - x_0) = P(0) = 0$, so we obtain

$$\begin{aligned}\operatorname{Bias}\hat{m}_n(x;p) &= \frac{1}{n}w(x_0,x_i)\sum_{i=1}^n w(x_0,x_i)\frac{m^{(\beta_0)}(x_0+\tau_i(x_i-x_0))}{\beta_0!}(x_i-x_0)^{\beta_0} \\ &= \frac{1}{n}w(x_0,x_i)\sum_{i=1}^n w(x_0,x_i)\frac{m^{(\beta_0)}(x_0+\tau_i(x_i-x_0)) - m^{(\beta_0)}(x_0)}{\beta_0!}(x_i-x_0)^{\beta_0},\end{aligned}$$

where the last line follows again from the reproducing property. Now we apply the fact that $m \in \mathcal{H}(\beta, L)$ with the previous lemma and find

$$\begin{aligned}|\operatorname{Bias}\hat{m}_n(x;p)| &\leq \frac{L}{n}\sum_{i=1}^n |w(x_0,x_i)|\frac{|x_i-x_0|^\beta}{\beta_0!} = \frac{L}{n}\sum_{i=1}^n |w(x_0,x_i)|\frac{|x_i-x_0|^\beta}{\beta_0!}\mathbb{1}_{|x_i-x|\leq h} \\ &\leq \frac{Lh^\beta}{n\beta_0!}\sum_{i=1}^n |w(x_0,x_i)| \leq \frac{8L\|K\|_\infty}{\lambda_0\beta_0!}h^\beta.\end{aligned}$$

$\square$

So again, we have a variance term of order $1/nh$ and a bias term of order h^β. However, both our bounds depend on λ_0, which depends on both n and x.

Proposition 3.6. *Suppose $x_i = i/n$ and that $K(u) \geq K_0 \mathbb{1}_{|u|\leq\Delta}$ for some $K_0, \Delta > 0$. Then, for $n \geq 2$ and $h \leq \frac{1}{4\Delta}$,*

$$\inf_{x\in[0,1]} \lambda_{0,n,x} \geq K_0 \inf_{v\in S^p} \min\left\{\int_0^\Delta (v^\top Q(u))^2 \,\mathrm{d}u\,, \int_{-\Delta}^0 (v^\top Q(u))^2 \,\mathrm{d}u\right\} - \frac{(4\Delta+2)K_0 e^{\Delta^2}}{nh}.$$

Proof. Since $\lambda_{0,n,x} = \inf_{v\in S^p} v^\top\big(\frac{1}{n}X^\top W X\big)v$, we will try to bound this quantity from below.

Let $u_i := \frac{x_i - x}{h}$ for $i = 1, \ldots, n$, so $u_1 \leq \frac{x_1}{h} = \frac{1}{nh}$, and let $u_0 := 0$.

First, we assume $x < 1 - h\Delta$, so that $u_n > \frac{1-(1-h\Delta)}{h} = \Delta$. Then, for any $v \in S^p$, we have

$$\begin{aligned}
v^\top\left(\frac{1}{n}X^\top W X\right) &= \frac{1}{n}(Xv)^\top W(Xv) = \frac{1}{n}\sum_{i=1}^n (Xv)_i^2 W_{ii} \\
&= \frac{1}{nh}\sum_{i=1}^n (v^\top Q(u_i))^2 K(u_i) \geq \frac{K_0}{nh}\sum_{i=1}^n (v^\top Q(u_i))^2 \mathbb{1}_{u_i\in[0,\Delta]}.
\end{aligned}$$

This is a Riemann sum (since $u_{i+1} - u_i = 1/(nh)$) and we will try to approximate it by the corresponding integral $K_0 \int_0^\Delta (v^\top Q(u))^2 \,\mathrm{d}u$.

To do this, note that we have for $u \in [0, \Delta]$ that

$$\|Q(u)\| \leq \left\{\sum_{\ell=0}^p \frac{u^{2\ell}}{(\ell!)}\right\}^{1/2} \leq \left(\sum_{\ell=0}^\infty \frac{u^{2\ell}}{\ell!}\right)^{1/2} \leq e^{\Delta^2/2}.$$

Furthermore, for $a, b \geq 0$, it can be shown that

$$\left|b^\ell - a^\ell\right| \leq \max(1, \ell/2)(b^{\ell-1} + a^{\ell-1})|b - a|,$$

and using this we obtain for $u, v \in [0, \Delta]$ that

$$\begin{aligned}
\|Q(u) - Q(v)\| &\leq \left\{\sum_{\ell=1}^p \frac{\max(1, \ell^2/4)(u^{\ell-1} + v^{\ell-1})^2(u-v)^2}{(\ell!)^2}\right\}^{1/2} \\
&\leq |u - v|\left\{\sum_{\ell=1}^p \frac{\max(1, \ell^2/4)(2\Delta^{\ell-1})^2}{(\ell!)^2}\right\}^{1/2} \leq |u - v|\left\{\sum_{\ell=1}^p \frac{\max(4, \ell^2)\Delta^{2\ell-2}}{(\ell!)^2}\right\}^{1/2} \\
&\overset{\star}{\leq} 2|u - v|\left\{\sum_{\ell=0}^{p-1} \frac{\Delta^{2\ell}}{(\ell!)^2}\right\}^{1/2} \leq 2|u - v|\left\{\sum_{\ell=0}^\infty \frac{\Delta^{2\ell}}{\ell!}\right\}^{1/2} = 2e^{\Delta^2/2}|u - v|,
\end{aligned}$$

where in $\star$ we used that $\max(4, \ell^2) \leq 4\ell^2$.

Now we estimate the difference between the Riemann sum and the corresponding integral. Write $i_- := \min\{i : u_i > 0\}$ and $i_+ := \min\{i \mid u_i > \Delta\}$. Then

$$\begin{aligned}
\frac{1}{nh}\sum_{i=1}^n (v^\top Q(u_i))^2 \mathbb{1}_{u_i\in[0,\Delta]} &= \sum_{i=i_-}^{i_+-1} \int_{u_{i-1}}^{u_i} (v^\top Q(u_i))^2 \,\mathrm{d}u \\
&\leq \sum_{i=i_-}^{i_+} \int_{\max(u_{i-1},0)}^{\min(u_i,\Delta)} (v^\top Q(u_i))^2 \,\mathrm{d}u + \frac{1}{nh}\sup_{u\in[0,\Delta]} (v^\top Q(u))^2.
\end{aligned}$$

(Note that the last integral, from u_{i-1} to Δ, is 0 since) Therefore we obtain

$$\begin{aligned}
&\left|\frac{1}{nh}\sum_{i=1}^{n}(v^\top Q(u_i))^2 \mathbb{1}_{u_i\in[0,\Delta]} - \int_0^\Delta (v^\top Q(u))^2 \,\mathrm{d}u\right| \\
&\leq \left|\sum_{i=i_-}^{i_+}\int_{\max(u_{i-1},0)}^{\min(u_i,\Delta)} (v^\top Q(u_i))^2 - (v^\top Q(u))^2 \,\mathrm{d}u\right| + \frac{1}{nh}\sup_{u\in[0,\Delta]}(v^\top Q(u))^2. \qquad (5)
\end{aligned}$$

Now note that by Cauchy-Schwarz since $\|v\| = 1$ we have

$$\begin{aligned}
\left|(v^\top Q(u_i))^2 - (v^\top Q(u))^2\right| &= \left|(v^\top Q(u_i) + v^\top Q(u))\right|\left|(v^\top Q(u_i) - v^\top Q(u))\right| \\
&\leq (\|Q(u_i)\| + \|Q(u)\|)\|Q(u_i) - Q(u)\| \\
&\leq 2e^{\Delta^2/2}\cdot 2e^{\Delta^2/2}|u_i - u| = 4e^{\Delta^2}|u_i - u|,
\end{aligned}$$

and therefore

$$(5) \leq 4e^{\Delta^2}\sum_{i=i_-}^{i_+}\int_{\max(u_{i-1},0)}^{\min(u_i,\Delta)}(u_i - u)\,\mathrm{d}u + \frac{e^{\Delta^2}}{nh} \leq \frac{(4\Delta+1)e^{\Delta^2}}{nh},$$

which concludes the case $x < 1 - h\Delta$.

Suppose $x \geq 1 - h\Delta$, then we have $u_1 \leq -\Delta$ and $u_n \geq 0$, and we can apply very similar arguments to reach the desired conclusion. □

Now that we have bounds on the variance and bias, we can prove our uniform bound:

Theorem 3.7. *Under the conditions of the previous two propositions, if we choose $h = \alpha n^{-1/(2\beta+1)}$ for some $\alpha > 0$, there exists $n_0 \in \mathbb{N}, C > 0$, depending only on $\beta, L, \alpha, K, \sigma^2_{\max}$, such that*

$$\sup_{m\in\mathcal{H}(\beta,L)}\sup_{x_0\in[0,1]}\mathbb{E}\Big[\{\hat{m}_n(x_0;p) - m(x_0)\}^2\Big] \leq Cn^{-2\beta/(2\beta+1)}.$$

3.3 Splines

3.3.1 Cubic splines

Let $n \geq 3$ and $a \leq x_1 < \cdots < x_n \leq b$.

Definition 3.8. A function $g\colon [a,b] \to \mathbb{R}$ is called a *cubic spline* with *knots* $x_1, \ldots, x_n$ if

1. g is a cubic polynomial on each interval $(a, x_1), (x_1, x_2), \ldots, (x_n, b)$;
2. $g \in C^2[a,b]$.

Furthermore, g is called *natural* if it is linear on $[a, x_1]$ and $[x_n, b]$, (i.e., $g''(a) = g'''(a) = g''(b) = g'''(b) = 0$).

We often represent a natural cubic spline by the vectors $\boldsymbol{g} \in \mathbb{R}^n$ with $g_i = g(x_i)$, and $\boldsymbol{\gamma} \in \mathbb{R}^{n-2}$ with $\gamma_i = g''(x_i)$ (excluding γ_1 and γ_n). Writing $h_i := x_{i+1} - x_i$ we have for $x \in [x_i, x_{i+1}]$ that

$$g(x) = \frac{(x - x_i)g_{i+1} - (x_{i+1} - x)g_i}{h_i} - \frac{1}{6}(x - x_i)(x_{i+1} - x)\left\{\left(1 + \frac{x - x_i}{h_i}\right)\gamma_{i+1} + \left(1 + \frac{x_{i+1} - x}{h_i}\gamma_i\right)\right\}.$$

Proposition 3.9. *Given $\boldsymbol{g} \in \mathbb{R}^n$, there exists a unique natural cubic spline g with knots at $x_1, \ldots, x_n$ satisfying $g(x_i) = g_i$ for all i, and there exists $K \succeq 0$ (depending on $x_1, \ldots, x_n$) such that*

$$\int_a^b g''(x)^2 \,\mathrm{d}x = \boldsymbol{g}^\top K\boldsymbol{g}.$$

We call g the *natural cubic spline interpolant to $\boldsymbol{g}$ at $x_1, \dots, x_n$*.

Definition 3.10. We define $\mathcal{S}_2[a,b]$ as the set of real-valued functions on $[a,b]$ with an absolutely continuous first derivative. For $f \in \mathcal{S}_2[a,b]$, we define the *roughness* of f by $R(f) \coloneqq \int_a^b f''(x)^2 \,\mathrm{d}x$.

Proposition 3.11. *For any $\boldsymbol{g} \in \mathbb{R}^n$, the natural cubic spline interpolant to $\boldsymbol{g}$ at $x_1, \dots, x_n$ is the unique minimiser of R over all $g \in \mathcal{S}_2[a,b]$ that satisfy $g(x_i) = g_i$ for all i.*

3.3.2 Natural cubic smoothing splines

Consider the nonparametric regression model $Y_i = g(x_i) + \sigma\varepsilon_i$, where $\mathbb{E}[\varepsilon_i] = 0$ and $\mathrm{Var}[\varepsilon_i] = 1$. A way to estimate a nonparametric regression function is to balance data fidelity against roughness of the curve, which can be done by minimising

$$S_\lambda(g) \coloneqq \sum_{i=1}^{n} (Y_i - g(x_i))^2 + \lambda R(g),$$

where $\lambda > 0$ is a regularisation parameter. For small λ, this is almost an exact fit to the data. For large λ, we are basically minimising $|g''|$, which means we will approximate the linear regression fit.

Theorem 3.12. *For each $\lambda \in (0, \infty)$, there exists a unique minimiser $\hat{g}_\lambda$ of S_λ over $\mathcal{S}_2[a,b]$. It is the natural cubic spline with knots at $x_1, \dots, x_n$ and $\boldsymbol{g} = (I + \lambda K)^{-1}\boldsymbol{Y}$.*

Proof. If g is not a natural cubic spline, we know that the natural cubic spline g^* which interpolates $g(x_1), \dots, g(x_n)$ at $x_1, \dots, x_n$ has a strictly lower value of S_λ, so we know the minimiser must be a natural cubic spline.

If g is a natural cubic spline, then there exists $K \succeq 0$ such that

$$\begin{aligned} S_\lambda(g) &= (\boldsymbol{Y} - \boldsymbol{g})^\top (\boldsymbol{Y} - \boldsymbol{g}) + \lambda \boldsymbol{g}^\top K \boldsymbol{g} \\ &= \boldsymbol{g}^\top (I + \lambda K)\boldsymbol{g} - 2\boldsymbol{Y}^\top \boldsymbol{g} + \boldsymbol{Y}^\top \boldsymbol{Y}, \end{aligned}$$

and by "completing the square" we write, for some Z independent of $\boldsymbol{g}$,

$$S_\lambda(g) = \big(\boldsymbol{g} - (I + \lambda K)^{-1}\boldsymbol{Y}\big)^\top (I + \lambda K)\big(\boldsymbol{g} - (I + \lambda K)^{-1}\boldsymbol{Y}\big) + Z$$

Since $I + \lambda K$ is positive definite it follows that $\boldsymbol{g} = (I + \lambda K)^{-1}\boldsymbol{Y}$ gives the minimiser. $\square$