Supplemental Material to
A SIMPLE ADJUSTMENT FOR BANDWIDTH SNOOPING

By

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A Simple Adjustment for Bandwidth Snooping: Supplemental Materials

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This supplement contains auxiliary results and proofs of the results in Section 5 of the main text, as well as tables of the critical values, descriptions of some of the details of our Monte Carlo study, and additional Monte Carlo simulations.

The following additional notation, which is also used in the appendix in the main text, is used throughout this supplement. For a sample $\{Z_i\}_{i=1}^n$ and a function f on the sample space, $E_n f(Z_i) = \frac{1}{n}\sum_{i=1}^n f(Z_i)$ denotes the sample mean, and $\mathbb{G}_n f(Z_i) = \sqrt{n}(E_n - E)f(Z_i) = \sqrt{n}[E_n f(Z_i) - Ef(Z_i)]$ denotes the empirical process. We use $t \vee t'$ and $t \wedge t'$ to denote element-wise maximum and minimum, respectively. We use e_k to denote the kth basis vector in Euclidean space (where the dimension of the space is clear from context).

S1 Auxiliary Results

This section contains auxiliary results that are used in the proof of Theorem 3.1 in Appendix A of the main text, and in the proofs of the results from Section 5 of the main text given later in this supplement.

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S1.1 Tail Bounds for Empirical Processes

We state some tail bounds based on an inequality of Talagrand (1996) and other empirical process results. Throughout this section, we consider a class of functions $\mathcal{G}$ on the sample space $\mathbb{R}^{d_Z}$ with an iid sample of random variables $Z_1, \ldots, Z_n$. We assume throughout that $\mathcal{G}$ has a polynomial covering number in the sense that, for some B, W, $N_1(\delta, Q, \mathcal{G}) \leq B\varepsilon^{-W}$ for all finitely discrete probability measures Q, where N_1 is defined in, e.g., Pollard (1984), p. 25.

Lemma S1.1. *Let $\tilde{\mathcal{G}}$ be a subset of $\mathcal{G}$ such that, for some envelope function G and constant $\overline{g}$, $|g(Z_i)| \leq G(Z_i) \leq \overline{g}$ a.s. for all $g \in \tilde{\mathcal{G}}$. Then, for some constant K that depends only on $\mathcal{G}$,*

$$P\left(\sup_{g\in\tilde{\mathcal{G}}} |\mathbb{G}_n g(Z_i)| \geq K\sqrt{E[G(Z_i)^2]} + t\right) \leq K \exp\left(-\frac{1}{K}\frac{t^2}{E[G(Z_i)^2] + \overline{g}\left\{\sqrt{E[G(Z_i)^2]} + t\right\}/\sqrt{n}}\right)$$

Proof. We apply a result of Talagrand (1996) as stated in equation (3) of Massart (2000). The quantity v from that version of the bound is, in our setting, given by $v = E\sup_{g\in\tilde{\mathcal{G}}}\sum_{i=1}^n [g(Z_i) - Eg(Z_i)]^2$ which, as shown in Massart (2000) (p.882), is bounded by $n\sup_{g\in\tilde{\mathcal{G}}} E\{[g(Z_i) - Eg(Z_i)]^2\} + 32\overline{g}E\sup_{g\in\tilde{\mathcal{G}}}\sum_{i=1}^n [g(Z_i) - Eg(Z_i)]$ (see also Klein and Rio, 2005). By Theorem 2.14.1 in van der Vaart and Wellner (1996),

$$E\sup_{g\in\tilde{\mathcal{G}}}\sum_{i=1}^n [g(Z_i) - Eg(Z_i)] \leq \sqrt{n}K_1\sqrt{E[G(Z_i)]^2}, \tag{1}$$

for a constant K_1 that depends only on $\mathcal{G}$. Combined with the fact that $E\{[g(Z_i) - Eg(Z_i)]^2\} \leq E[G(Z_i)^2]$, this gives the bound

$$v \leq nE[G(Z_i)^2] + 32\overline{g}K_1\sqrt{n}\sqrt{E[G(Z_i)]^2}.$$

Applying the bound from equation (3) of Massart (2000) with these quantities gives

$$
\begin{aligned}
&P\left(\sqrt{n}\sup_{g\in\tilde{\mathcal{G}}}\mathbb{G}_n g(Z_i)\geq K_1\sqrt{n}\sqrt{E[G(Z_i)]^2}+r\right)\\
&\leq P\left(\sqrt{n}\sup_{g\in\tilde{\mathcal{G}}}\mathbb{G}_n g(Z_i)\geq E\sup_{g\in\tilde{\mathcal{G}}}\sum_{i=1}^{n}[g(Z_i)-Eg(Z_i)]+r\right)\\
&\leq K_2\exp\left(-\frac{1}{K_2}\frac{r^2}{nE[G(Z_i)^2]+32\overline{g}K_1\sqrt{n}\sqrt{E[G(Z_i)]^2}+\overline{g}r}\right)
\end{aligned}
$$

where the first inequality follows from (1). Substituting $r=\sqrt{n}t$ gives

$$
\begin{aligned}
&P\left(\sup_{g\in\tilde{\mathcal{G}}}\mathbb{G}_n g(Z_i)\geq K_1\sqrt{E[G(Z_i)]^2}+t\right)\\
&\leq K_2\exp\left(-\frac{1}{K_2}\frac{t^2}{E[G(Z_i)^2]+32\overline{g}K_1\sqrt{E[G(Z_i)]^2}/\sqrt{n}+\overline{g}t/\sqrt{n}}\right)
\end{aligned}
$$

which gives the result after noting that replacing K_1 on the left hand side as well as K_2 and $32K_1K_2$ on the right hand side with a larger constant K decreases the left hand side and increases the right hand side, and applying a symmetric bound to $\inf_{g\in\tilde{\mathcal{G}}}\mathbb{G}_n g(Z_i)$. □

The above lemma gives good bounds for t just larger than $\sqrt{E[G(Z_i)]^2}$, so long as $\sqrt{E[G(Z_i)]^2}/\sqrt{n}$ is small relative to $E[G(Z_i)]^2$ (i.e. so long as $E[G(Z_i)]^2 n$ is large). We now state a version of this result that is specialized to this case.

Lemma S1.2. *Let $\tilde{\mathcal{G}}$ be a subset of $\mathcal{G}$ such that, for some envelope function G and constant $\overline{g}$, $|g(Z_i)|\leq G(Z_i)\leq\overline{g}$ a.s. for all $g\in\tilde{\mathcal{G}}$. Then, for some constant K that depends only on $\mathcal{G}$,*

$$
P\left(\sup_{g\in\tilde{\mathcal{G}}}|\mathbb{G}_n g(Z_i)|\geq\sqrt{V}a\right)\leq K\exp\left(-\frac{a^2}{K}\right)
$$

for all $V\geq E[G(Z_i)^2]$ and $a>0$ with $a+1\leq\sqrt{V}\sqrt{n}/\overline{g}$.

Proof. Substituting $t=rV^{1/2}$ into the bound from Lemma S1.1 gives, letting K_1 be the constant

K from that lemma,

$$P\left(\sup_{g\in\tilde{\mathcal{G}}}|\mathbb{G}_n g(Z_i)| \geq (K_1+r)V^{1/2}\right) \leq K_1 \exp\left(-\frac{1}{K_1}\frac{r^2 V}{V+\overline{g}\left\{V^{1/2}+rV^{1/2}\right\}/\sqrt{n}}\right).$$

For $\overline{g}(1+r) \leq \sqrt{n}V^{1/2}$, this is bounded by $K_1 \exp\left(-\frac{r^2}{2K_1}\right)$. Setting $a = K_1 + r$ and noting that $K_1 \exp\left(-\frac{(a-K_1)^2}{2K_1}\right) \leq K_2 \exp\left(-\frac{a^2}{K_2}\right)$ for a large enough constant K_2 (and that $\overline{g}(1+a) \leq \sqrt{n}V^{1/2}$ implies $\overline{g}(1+a-K_1) \leq \sqrt{n}V^{1/2}$) gives the result. □

S1.2 Tail Bounds for Kernel Estimators

We specialize some of the results of Section S1.1 to our setting. We are interested in functions of the form $g(x,w) = f(w,h,t)k(x/h)$, where h varies over positive real numbers and t varies over some index set T.

We assume throughout the section that $k(x)$ is a bounded kernel function with support $[-A, A]$, with $k(x) \leq B_k < \infty$ for all k. We also assume that X_i is a real valued random variable with with a density $f_X(x)$ with $f_X(x) \leq \overline{f}_X < \infty$ all x.

Lemma S1.3. *Suppose that* $\{(x,w) \mapsto f(w,h,t)k(x/h) | 0 \leq h \leq \overline{h}, t \in T\}$ *is contained in some larger class* $\mathcal{G}$ *with polynomial covering number, and that, for some constant* B_f, $|f(W_i,h,t)k(X_i/h)| \leq B_f$ *for all* $h \leq \overline{h}$ *and* $t \in T$ *with probability one. Then, for some constant* K *that depends only on* $\mathcal{G}$,

$$P\left(\sup_{0\leq h\leq \overline{h}, t\in T}|\mathbb{G}_n f(W_i,h,t)k(X_i/h)| \geq aB_f A^{1/2}\overline{f}_X^{1/2}\overline{h}^{1/2}\right) \leq K\exp(-\frac{a^2}{K})$$

for all $a > 0$ *with* $a + 1 \leq A^{1/2}\overline{f}_X^{-1/2}\overline{h}^{-1/2}n^{1/2}$.

Proof. The result follows from Lemma S1.2, since $B_f I(|X_i| \leq A\overline{h})$ is an envelope function for $f(W_i,h,t)k(X_i/h)$ as h and t vary over this set. □

Lemma S1.4. *Suppose that the conditions of Lemma S1.3 hold, and let* $a(h) = 2\sqrt{K \log\log(1/h)}$ *where*

K is the constant from Lemma S1.3. Then, for a constant $\varepsilon > 0$ that depends only on K, A and $\overline{f}_X$,

$$P\left(|\mathbb{G}_n f(W_i,h,t)k(X_i/h)| \geq a(h)h^{1/2}B_f A^{1/2}\overline{f}_X^{1/2} \text{ some } (\log\log n)/(\varepsilon n) \leq h \leq \overline{h},\, t \in T\right)$$
$$\leq K(\log 2)^{-2} \sum_{(2\overline{h})^{-1} \leq 2^k \leq \infty} k^{-2}.$$

Proof. Let $\mathcal{H}^k = (2^{-(k+1)}, 2^{-k})$. Applying Lemma S1.3 to this set, we have

$$P\left(|\mathbb{G}_n f(W_i,h,t)k(X_i/h)| \geq a(h)h^{1/2}B_f A^{1/2}\overline{f}_X^{1/2} \text{ some } h \in \mathcal{H}^k,\, t \in T\right)$$
$$\leq P\left(\sup_{0 \leq h \leq 2^k, t \in T} |\mathbb{G}_n f(W_i,h,t)k(X_i/h)| \geq a(2^{-k})2^{-(k+1)/2}B_f A^{1/2}\overline{f}_X^{1/2}\right)$$
$$\leq K\exp\left(-\frac{[a(2^{-k})2^{-1/2}]^2}{K}\right) = K\exp\left(-2\log\log 2^k\right) = K\exp\left(-2\log(k\log 2)\right) = K[k\log 2]^{-2}$$

so long as $2^{-1/2}a(2^{-k}) + 1 \leq A^{1/2}\overline{f}_X^{1/2}2^{-k/2}n^{1/2}$, where the first inequality follows since $a(h) \geq a(2^{-k})$ and $h \geq 2^{-(k+1)}$ for $h \in \mathcal{H}^k$.

Note that $2^{-1/2}a(2^{-k}) + 1 \leq A^{1/2}\overline{f}_X^{1/2}2^{-k/2}n^{1/2}$ will hold iff. $[2^{-1/2}a(2^{-k}) + 1]2^{k/2} \leq A^{1/2}\overline{f}_X^{1/2}n^{1/2}$. If $2^k \leq \varepsilon n/\log\log n$ for some $\varepsilon > 0$, we will have $a(2^{-k}) \leq 2\sqrt{K\log\log[\varepsilon n/\log\log n]}$, so that $[2^{-1/2}a(2^{-k}) + 1]2^{k/2} \leq \{2^{-1/2} \cdot 2\sqrt{K\log\log[\varepsilon n/\log\log n]} + 1\}\sqrt{\varepsilon n/\log\log n}$. For large enough n, this is bounded by $4\sqrt{K\varepsilon n}$, which is less than $A^{1/2}\overline{f}_X^{1/2}n^{1/2}$ for ε small enough as required.

Thus, for ε defined above,

$$P\left(|\mathbb{G}_n f(W_i,h,t)k(X_i/h)| \geq a(h)h^{1/2}B_f A^{1/2}\overline{f}_X^{1/2} \text{ some } (\log\log n)/(\varepsilon n) \leq h \leq \overline{h},\, t \in T\right)$$
$$\leq \sum_{(2\overline{h})^{-1} \leq 2^k \leq 2\varepsilon n/\log\log n} P\left(\sup_{0 \leq h \leq 2^k, t \in T} |\mathbb{G}_n f(W_i,h,t)k(X_i/h)| \geq a(2^{-k})2^{-(k+1)/2}B_f A^{1/2}\overline{f}_X^{1/2}\right)$$
$$\leq K(\log 2)^{-2} \sum_{(2\overline{h})^{-1} \leq 2^k \leq 2\varepsilon n/\log\log n} k^{-2},$$

which gives the result. □

Using these bounds, we obtain the following uniform bound on $\mathbb{G}_n f(W_i,h,t)k(X_i/h)$.

Lemma S1.5. *Under the conditions of Lemma S1.4,*

$$\sup_{(\log\log n)/(\varepsilon n)\le h\le \overline{h}, t\in T} \frac{|\mathbb{G}_n f(W_i,h,t)k(X_i/h)|}{(\log\log h^{-1})^{1/2}h^{1/2}} = \mathcal{O}_P(1)$$

Proof. Given $\varepsilon > 0$, we can apply Lemma S1.4 to find a $\delta > 0$ such that

$$\sup_{(\log\log n)/(\varepsilon n)\le h\le \delta, t\in T} \frac{|\mathbb{G}_n f(W_i,h,t)k(X_i/h)|}{(\log\log h^{-1})^{1/2}h^{1/2}} < 2\sqrt{2K}B_f A^{1/2}\overline{f}_X^{1/2}$$

with probability at least $1 - K(\log 2)^{-2}\sum_{(2\delta)^{-1}\le 2^k\le\infty} k^{-2} > 1-\varepsilon/2$. For this choice of δ,

$$\sup_{\delta\le h\le \overline{h}, t\in T} \frac{|\mathbb{G}_n f(W_i,h,t)k(X_i/h)|}{(\log\log h^{-1})^{1/2}h^{1/2}} = \mathcal{O}_P(1)$$

by Lemma S1.3. Thus, choosing C large enough so that $C \ge 2\sqrt{2K}B_f A^{1/2}\overline{f}_X^{1/2}$ and

$$\sup_{\delta\le h\le \overline{h}, t\in T} \frac{|\mathbb{G}_n f(W_i,h,t)k(X_i/h)|}{(\log\log h^{-1})^{1/2}h^{1/2}} \le C$$

with probability at least $1-\varepsilon/2$ asymptotically, we have $\sup_{(\log\log n)/(\varepsilon n)\le h\le \overline{h}, t\in T} \frac{|\mathbb{G}_n f(W_i,h,t)k(X_i/h)|}{(\log\log h^{-1})^{1/2}h^{1/2}} \le C$ with probability at least $1-\varepsilon$ asymptotically. □

S1.3 Gaussian Approximation

This section proves Theorem A.2 in Appendix A.4, which gives a Gaussian process approximation for the process $\hat{\mathbb{H}}_n(h)$ defined in that section.

For convenience, we repeat the setup here. We show that $\frac{1}{\sqrt{h}}\mathbb{G}_n \tilde{Y}_i k(X_i/h) = \frac{1}{\sqrt{nh}}\sum_{i=1}^n \tilde{Y}_i k(X_i/h)$ is approximated by a Gaussian process with the same covariance kernel. We consider a general setup with $\{(\tilde{X}_i,\tilde{Y}_i)\}_{i=1}^n$ iid, with $\tilde{X}_i \ge 0$ a.s. such that $\tilde{X}_i$ has a density $f_{\tilde{X}}(x)$ on $[0,\overline{x}]$ for some $\overline{x}\ge 0$, with $f_{\tilde{X}}(x)$ bounded away from zero and infinity on this set. We assume that $\tilde{Y}_i$ is bounded almost surely, with $E(\tilde{Y}_i|\tilde{X}_i) = 0$ and $var(\tilde{Y}_i|\tilde{X}_i = x) = f_{\tilde{X}}(x)^{-1}$. We assume that the kernel function k has finite support $[0,A]$ and is differentiable on its support with bounded derivative. For ease of notation, we assume in this section that $\int k(u)^2\,du = 1$. The result applies to our setup with $\tilde{Y}_i$ given in (11) in Section A of the appendix in the main text and $\tilde{X}_i$ given by $|X_i|$.

Let

$$\hat{\mathbb{H}}_n(h) = \frac{1}{\sqrt{nh}} \sum_{i=1}^{n} \tilde{Y}_i k(\tilde{X}_i/h).$$

Theorem A.2. Under the conditions above, there exists, for each n, a process $\mathbb{H}_n(h)$ such that, conditional on $(\tilde{X}_1, \ldots, \tilde{X}_n)$, $\mathbb{H}_n$ is a Gaussian process with covariance kernel

$$cov\left(\mathbb{H}_n(h), \mathbb{H}_n(h')\right) = \frac{1}{\sqrt{hh'}} \int k(x/h)k(x/h')\, dx$$

and

$$\sup_{\underline{h}_n \le h \le \overline{x}/A} \left|\hat{\mathbb{H}}_n(h) - \mathbb{H}_n(h)\right| = \mathcal{O}_P\left((n\underline{h}_n)^{-1/4}[\log(n\underline{h}_n)]^{1/2}\right)$$

for any sequence $\underline{h}_n$ with $n\underline{h}_n / \log\log \underline{h}_n^{-1} \to \infty$.

We now prove the result. Let $\hat{G}(x) = \frac{1}{n}\sum_{\tilde{X}_i \le x} \tilde{Y}_i$. With this notation, we can write the process $\hat{\mathbb{H}}_n(h)$ as

$$\hat{\mathbb{H}}_n(h) = \frac{1}{\sqrt{nh}} \sum_{i=1}^{n} \tilde{Y}_i k(\tilde{X}_i/h) = \frac{\sqrt{n}}{\sqrt{h}} \int k(x/h)\, d\hat{G}(x).$$

Let $\hat{g}(x) = \frac{1}{n}\sum_{\tilde{X}_i \le x} f_{\tilde{X}}(\tilde{X}_i)^{-1}$. In Lemma S1.6 below, a process $\mathbb{B}_n(t)$ is constructed that is a Brownian motion conditional on $\tilde{X}_1, \ldots, \tilde{X}_n$ such that $\mathbb{B}_n(n\hat{g}(x))$ is, with high probability conditional on $\tilde{X}_1, \ldots, \tilde{X}_n$, close to $n\hat{G}(x)$. By showing that $\hat{g}(x)$ is close to x with high probability and using properties of the fluctuation of the Brownian motion, it is then shown that $\mathbb{B}_n(n\hat{g}(x))$ can be approximated by $\mathbb{B}_n(nx)$, so that $\hat{\mathbb{H}}_n(h)$ is approximated by the corresponding process with $\hat{G}(x)$ replaced by $\mathbb{B}_n(nx)/n$.

Formally, let $\mathbb{B}_n(t)$ be given by the (conditional) Brownian motion in Lemma S1.6 below, and define

$$\mathbb{H}_n(h) = \frac{1}{\sqrt{nh}} \int k(x/h)\, d\mathbb{B}_n(nx).$$

Note that $\mathbb{H}_n(h) = \frac{1}{\sqrt{h}} \int k(x/h)\, d\tilde{\mathbb{B}}_n(x)$ (where $\tilde{\mathbb{B}}_n(x) = \mathbb{B}_n(nx)/\sqrt{n}$ is another Brownian motion conditional on $\tilde{X}_1, \ldots, \tilde{X}_n$), so that, conditional on $(\tilde{X}_1, \ldots, \tilde{X}_n)$, $\mathbb{H}_n$ is a Gaussian process with the desired covariance kernel.

Let $R_{1,n}(x) = n\hat{G}(x) - \mathbb{B}_n(n\hat{g}(x))$ and $R_{2,n}(x) = \mathbb{B}_n(n\hat{g}(x)) - \mathbb{B}_n(nx)$. Then

$$\hat{\mathbb{H}}_n(h) - \mathbb{H}_n(h) = \frac{1}{\sqrt{nh}} \int k(x/h)\, dR_{1,n}(x) + \frac{1}{\sqrt{nh}} \int k(x/h)\, dR_{2,n}(x).$$

Using the integration by parts formula, we have, for $j = 1, 2$ and $Ah \le \overline{x}$,

$$\frac{1}{\sqrt{nh}} \int k(x/h)\, dR_{j,n}(x) = \frac{R_{j,n}(Ah)k(A)}{\sqrt{nh}} - \frac{1}{\sqrt{nh}} \int_{x=0}^{Ah} R_{j,n}(x) k'(x/h) \frac{1}{h}\, dx$$

The first term is bounded by $\frac{|R_{j,n}(Ah)|k(A)}{\sqrt{nh}}$, and the second term is bounded by

$$\frac{A}{\sqrt{nh}} \left(\sup_{0 \le x \le Ah} |R_{j,n}(x)| \right) \left(\sup_{0 \le u \le A} |k'(u)| \right)$$

(see Bickel and Rosenblatt, 1973, for a similar derivation). By boundedness of $k'(u)$, it follows that both terms are bounded by a constant times $\frac{1}{\sqrt{nh}} \sup_{0 \le x \le Ah} |R_{j,n}(x)|$, so that

$$\sup_{\underline{h}_n \le h \le \overline{x}/A} |\hat{\mathbb{H}}_n(h) - \mathbb{H}_n(h)| \le K \sup_{\underline{h}_n \le h \le \overline{x}/A} \sum_{j=1}^{2} \sup_{0 \le x \le Ah} \frac{|R_{j,n}(x)|}{\sqrt{nh}} \le K \sum_{j=1}^{2} \sup_{0 \le x \le \overline{x}} \frac{|R_{j,n}(x)|}{\sqrt{n[(x/A) \vee \underline{h}_n]}}.$$

for some constant K. Thus, the result will follow if we can show that $\sup_{0 \le x \le \overline{x}} \frac{|R_{1,n}(x)|}{\sqrt{n(x \vee \underline{h}_n)}}$ and $\sup_{0 \le x \le \overline{x}} \frac{|R_{2,n}(x)|}{\sqrt{n(x \vee \underline{h}_n)}}$ converge to zero at the required rate.

We first construct $\mathbb{B}_n(t)$ and show that $\sup_{0 \le x \le A\overline{x}/A} \frac{|R_{1,n}(x)|}{\sqrt{n(x \vee \underline{h}_n)}}$ converges to zero quickly enough with this construction, using an approximation of Sakhanenko. Let $\hat{F}_{\tilde{X}}(x) = \frac{1}{n} \sum_{i=1}^{n} I(\tilde{X}_i \le x)$ be the empirical cdf of $\tilde{X}_i$, and let $\tilde{X}_{(k)}$ be the kth smallest value of $\tilde{X}_i$.

Lemma S1.6. *Under the conditions of Theorem A.2, one can construct variables $Z_1, \dots, Z_n$ with $Z_i | (\tilde{X}_1, \dots, \tilde{X}_n) \sim N(0, f_{\tilde{X}}(\tilde{X}_i)^{-1})$ such that*

$$P\left(\left| \sum_{\tilde{X}_i \le x} Z_i - \sum_{\tilde{X}_i \le x} \tilde{Y}_i \right| > K \log\left[n\hat{F}_{\tilde{X}}(x) + 2 \right] \text{ some } 0 \le x \le \overline{x} \,\middle|\, \tilde{X}_1, \dots, \tilde{X}_n \right) \le \varepsilon(K)$$

with probability one, where $\varepsilon(K)$ is a deterministic function with $\varepsilon(K) \to 0$ as $K \to \infty$.

Proof. Using a result of Sakhanenko (1985) as stated in Theorem A of Shao (1995), we can construct $Z_1, \dots, Z_n$ such that

$$E \exp\left(\lambda A \sup_{0 \le x \le \tilde{X}_{(k)}} \left| \sum_{\tilde{X}_i \le x} Z_i - \sum_{\tilde{X}_i \le x} \tilde{Y}_i \right| \middle| \tilde{X}_1, \dots, \tilde{X}_n \right) \le 1 + \lambda \sum_{\tilde{X}_i \le \tilde{X}_{(k)}} f_{\tilde{X}}(\tilde{X}_i)^{-1}$$

where A is a universal constant and λ is any constant such that $\lambda E[\exp(\lambda|\tilde{Y}_i|)|\tilde{Y}_i|^3|\tilde{X}_i] \le E[\tilde{Y}_i^2|\tilde{X}_i]$. Let $\overline{Y}$ be a bound for $\tilde{Y}_i$. Then $\lambda E[\exp(\lambda|\tilde{Y}_i|)|\tilde{Y}_i|^3|\tilde{X}_i] \le \lambda \exp(\lambda\overline{Y})\overline{Y} E[|\tilde{Y}_i|^2|\tilde{X}_i]$, so the inequality holds for any λ with $\lambda \exp(\lambda\overline{Y})\overline{Y} \le 1$. From now on, we fix $\lambda > 0$ so that this inequality holds.

Letting $\underline{f}_{\tilde{X}}$ be a lower bound for $f_{\tilde{X}}(x)$ over $0 \le x \le \overline{x}$ and applying Markov's inequality, the above bound gives

$$\begin{aligned}
&P\left(\lambda A \sup_{0 \le x \le \tilde{X}_{(k)}} \left| \sum_{\tilde{X}_i \le x} Z_i - \sum_{\tilde{X}_i \le x} \tilde{Y}_i \right| > t \middle| \tilde{X}_1, \dots, \tilde{X}_n \right) \\
&\le \exp(-t) E \exp\left(\lambda A \sup_{0 \le x \le \tilde{X}_{(k)}} \left| \sum_{\tilde{X}_i \le x} Z_i - \sum_{\tilde{X}_i \le x} \tilde{Y}_i \right| \middle| \tilde{X}_1, \dots, \tilde{X}_n \right) \le \exp(-t)(1 + \lambda \underline{f}_{\tilde{X}}^{-1} k).
\end{aligned}$$

Thus,

$$\begin{aligned}
&P\left(\left| \sum_{\tilde{X}_i \le x} Z_i - \sum_{\tilde{X}_i \le x} \tilde{Y}_i \right| > K \log\left[\sum_{i=1}^n I(\tilde{X}_i \le x) + 2\right] \text{ some } 0 \le x \le \overline{x} \middle| \tilde{X}_1, \dots, \tilde{X}_n \right) \\
&\le P\left(\sup_{0 \le x \le \tilde{X}_{(k)}} \left| \sum_{\tilde{X}_i \le x} Z_i - \sum_{\tilde{X}_i \le x} \tilde{Y}_i \right| > K \log k \text{ some } 2 \le k \le n \middle| \tilde{X}_1, \dots, \tilde{X}_n \right) \\
&\le \sum_{k=2}^n P\left(\lambda A \sup_{0 \le x \le \tilde{X}_{(k)}} \left| \sum_{\tilde{X}_i \le x} Z_i - \sum_{\tilde{X}_i \le x} \tilde{Y}_i \right| \ge \lambda A K \log k \middle| \tilde{X}_1, \dots, \tilde{X}_n \right) \\
&\le \sum_{k=2}^n k^{-\lambda A K}(1 + \lambda \underline{f}_{\tilde{X}}^{-1} k) \le \sum_{k=2}^\infty k^{-\lambda A K}(1 + \lambda \underline{f}_{\tilde{X}}^{-1} k),
\end{aligned}$$

which can be made arbitrarily small by making K large.

□

Embedding $\sum_{\tilde{X}_i \le x} Z_i$ in a Brownian motion, we can restate the above construction as follows:

with probability at least $1 - K(\varepsilon)$ conditional on $\tilde{X}_1, \ldots, \tilde{X}_n$,

$$\left| n\hat{G}(x) - \mathbb{B}_n(n\hat{g}(x)) \right| \leq K \log[n\hat{F}_{\tilde{X}}(x) + 2] \text{ all } 0 \leq x \leq \overline{x}$$

where $\mathbb{B}_n(t) = \mathbb{B}_n(t; \tilde{X}_1, \ldots, \tilde{X}_n)$ is a Brownian motion conditional on $\tilde{X}_1, \ldots, \tilde{X}_n$. Let $\overline{f}_{\tilde{X}}$ be an upper bound for the density of $\tilde{X}_i$ on $[0, \overline{x}]$.

Lemma S1.7. *Under the conditions of Theorem A.2, for any $\eta > 0$,*

$$\hat{F}_{\tilde{X}}(x) \leq \overline{f}_{\tilde{X}} \cdot (1 + \eta)(x \vee \underline{h}_n)$$

for all $0 \leq x \leq \overline{x}$ with probability approaching one.

Proof. By Lemma S1.5,

$$\sup_{\underline{h}_n \leq x \leq \overline{x}} \frac{\sqrt{n}|\hat{F}_{\tilde{X}}(x) - F_{\tilde{X}}(x)|}{\sqrt{x \log\log x^{-1}}} = \mathcal{O}_P(1).$$

Thus,

$$\begin{aligned} \sup_{\underline{h}_n \leq x \leq \overline{x}} \frac{|\hat{F}_{\tilde{X}}(x) - F_{\tilde{X}}(x)|}{x} &= \sup_{\underline{h}_n \leq x \leq \overline{x}} \frac{\sqrt{n}|\hat{F}_{\tilde{X}}(x) - F_{\tilde{X}}(x)|}{\sqrt{x \log\log x^{-1}}} \frac{\sqrt{x \log\log x^{-1}}}{\sqrt{n}x} \\ &= \mathcal{O}_P\left(\sup_{\underline{h}_n \leq x \leq \overline{x}} \frac{\sqrt{\log\log x^{-1}}}{\sqrt{nx}} \right) = \mathcal{O}_P\left(\frac{\sqrt{\log\log \underline{h}_n^{-1}}}{\sqrt{n\underline{h}_n}} \right) = o_P(1) \end{aligned}$$

where the last step follows since $n\underline{h}_n / \log\log \underline{h}_n^{-1} \to \infty$. Thus, for any $\eta > 0$, we have, with probability approaching one,

$$\hat{F}_{\tilde{X}}(x) \leq \hat{F}_{\tilde{X}}(x \vee \underline{h}_n) \leq F_{\tilde{X}}(x \vee \underline{h}_n) + (\eta \overline{f}_{\tilde{X}})(x \vee \underline{h}_n) \leq \overline{f}_{\tilde{X}} \cdot (1 + \eta)(x \vee \underline{h}_n)$$

for all x. □

Combining these two lemmas, we have, for large enough n,

$$\limsup_n P\left(\left|n\hat{G}(x) - \mathbb{B}_n(n\hat{g}(x))\right| > K\log\left[2n\overline{f}_{\tilde{X}}(x \vee \underline{h}_n) + 2\right] \text{ some } 0 \le x \le \overline{x}\right)$$
$$\le \varepsilon(K) + \limsup_n P\left(\hat{F}_{\tilde{X}}(x) > \overline{f}_{\tilde{X}} \cdot 2(x \vee \underline{h}_n)\right) \le \varepsilon(K).$$

Since this can be made arbitrarily small by making K large, it follows that

$$\sup_{0\le x\le\overline{x}} \frac{\left|n\hat{G}(x) - \mathbb{B}_n(n\hat{g}(x))\right|}{\sqrt{n(x \vee \underline{h}_n)}} = \mathcal{O}_P\left(\sup_{0\le x\le\overline{x}} \frac{\log\left[2n\overline{f}_{\tilde{X}}(x \vee \underline{h}_n) + 2\right]}{\sqrt{n(x \vee \underline{h}_n)}}\right) = \mathcal{O}_P\left(\frac{\log(n\underline{h}_n)}{\sqrt{n\underline{h}_n}}\right),$$

which gives the required rate for $R_{1,n}(x)$.

Define the function $LL(x) = \log\log x$ for $\log\log x \ge 1$ and $LL(x) = 1$ otherwise. Given K, let $B_n(K)$ be the event that

$$|n\hat{g}(x) - nx| \le K\sqrt{n(x \vee \underline{h}_n)LL(x/\underline{h}_n)} \text{ all } 0 \le x \le \overline{x},$$

and let $C_n(K)$ be the event that

$$\left|\mathbb{B}_n(t') - \mathbb{B}_n(t)\right| \le K\sqrt{(|t'-t| \vee 1)\cdot\log(t \vee t' \vee 2)} \text{ all } 0 \le t, t' < \infty.$$

Lemma S1.8. *On the event $B_n(K) \cap C_n(K)$, for large enough n,*

$$\frac{|R_{2,n}(x)|}{\sqrt{n(x \vee \underline{h}_n)}} \le K^{3/2}[n(x \vee \underline{h}_n)]^{-1/4}\{LL(x/\underline{h}_n)\}^{1/4} \cdot \{\log 2 + \log[n(x \vee \underline{h}_n)]\}^{1/2}$$
$$\le K^{3/2}(n\underline{h}_n)^{-1/4} \cdot \{\log 2 + \log[n\underline{h}_n]\}^{1/2}$$

for all $0 \le x \le \overline{x}$.

Proof. On this event, for all $0 \le x \le \overline{x}$ and large enough n,

$$
\begin{aligned}
&|R_{2,n}(x)| = |\mathbb{B}_n(n\hat{g}(x)) - \mathbb{B}_n(nx)| \le \sup_{|t-nx|\le K\sqrt{n(x\vee \underline{h}_n)LL(x/\underline{h}_n)}} |\mathbb{B}_n(t) - \mathbb{B}_n(nx)| \\
&\le \sup_{|t-nx|\le K\sqrt{n(x\vee \underline{h}_n)LL(x/\underline{h}_n)}} K\sqrt{(|t-nx|\vee 1)\cdot \log[t\vee (nx)\vee 2]} \\
&\le K\sqrt{K\sqrt{n(x\vee \underline{h}_n)LL(x/\underline{h}_n)}\cdot \log[2n(x\vee \underline{h}_n)]} \\
&= K^{3/2}n^{1/4}(x\vee \underline{h}_n)^{1/4}\{LL(x/\underline{h}_n)\}^{1/4}\cdot \{\log 2 + \log[n(x\vee \underline{h}_n)]\}^{1/2}.
\end{aligned}
$$

□

Lemma S1.9. *Under the conditions of Theorem A.2, for any $\varepsilon > 0$, there exists a K such that $P(B_n(K)) \ge 1-\varepsilon$ for large enough n.*

Proof. Let $\mathcal{X}^k = (2^k\underline{h}_n, 2^{k+1}\underline{h}_n] \cap [0, \overline{x}]$. We have, for $k \ge 2$,

$$
\begin{aligned}
&P\left(|n\hat{g}(x) - nx| > K\sqrt{n(x\vee \underline{h}_n)LL(x/\underline{h}_n)} \text{ some } x \in \mathcal{X}^k\right) \\
&= P\left(|\mathbb{G}_n f(\tilde{X}_i)^{-1}I(\tilde{X}_i \le x)| > K\sqrt{x\cdot LL(x/\underline{h}_n)} \text{ some } x \in \mathcal{X}^k\right) \\
&\le P\left(\sup_{x\in\mathcal{X}^k} |\mathbb{G}_n f(\tilde{X}_i)^{-1}I(\tilde{X}_i \le x)| > K\sqrt{2^k\underline{h}_n\cdot LL(2^k)}\right) \\
&\le C\exp\left(-\frac{K^2 LL(2^k)}{C}\right) \le C\exp\left(-\frac{K^2}{C}\log\log(2^k)\right) = C[k\log 2]^{-\frac{K^2}{C}}
\end{aligned}
$$

for some constant C by Lemma S1.3. Thus,

$$
P\left(|n\hat{g}(x) - nx| > K\sqrt{n(x\vee \underline{h}_n)LL(x/\underline{h}_n)} \text{ some } 4\underline{h}_n \le x \le \overline{x}\right) \le C\sum_{k=2}^{\infty}[k\log 2]^{-K^2/C}
$$

which can be made arbitrarily small by making K large. Note also that

$$
\begin{aligned}
&P\left(|n\hat{g}(x) - nx| > K\sqrt{n(x\vee \underline{h}_n)LL(x/\underline{h}_n)} \text{ some } 0 \le x \le 4\underline{h}_n\right) \\
&\le P\left(\sup_{0\le x\le 4\underline{h}_n} |\mathbb{G}_n f(\tilde{X}_i)^{-1}I(\tilde{X}_i \le x)| > K\sqrt{\underline{h}_n}\right),
\end{aligned}
$$

which can also be made arbitrarily small by choosing K large by Lemma S1.3. Combining these bounds gives the result. □

Lemma S1.10. *Under the conditions of Theorem A.2, for any $\varepsilon > 0$, there exists a K such that $P(C_n(K)|\tilde{X}_1, \dots, \tilde{X}_n) \geq 1 - \varepsilon$ with probability one for all n.*

Proof. We have

$$
\begin{aligned}
&1 - P(C_n(K)|\tilde{X}_1, \dots, \tilde{X}_n) = P\left(\left|\mathbb{B}_n(t') - \mathbb{B}_n(t)\right| > K\sqrt{(|t-t'| \vee 1) \cdot \log(t \vee t' \vee 2)} \text{ some } 0 \leq t, t' < \infty \right) \\
&= P\left(|\mathbb{B}_n(t+s) - \mathbb{B}_n(t)| > K\sqrt{(s \vee 1) \cdot \log[(t+s) \vee 2]} \text{ some } 0 \leq s, t < \infty \right) \\
&\leq \sum_{k=0}^{\infty} \sum_{\ell=0}^{\infty} P\left(|\mathbb{B}_n(t+s) - \mathbb{B}_n(t)| > K\sqrt{(s \vee 1) \cdot \log[(t+s) \vee 2]} \text{ some } (s,t) \in \mathcal{S}_{k,\ell} \right)
\end{aligned}
$$

where $\mathcal{S}_{k,\ell} = \{(s,t) | \ell \leq s \leq \ell + 1, (\ell \vee 1)k \leq t \leq (\ell \vee 1)(k+1)\}$. Note that

$$
\begin{aligned}
&P\left(|\mathbb{B}_n(t+s) - \mathbb{B}_n(t)| > K\sqrt{(s \vee 1) \cdot \log[(t+s) \vee 2]} \text{ some } (s,t) \in \mathcal{S}_{k,\ell} \right) \\
&\leq P\left(|\mathbb{B}_n(t+s) - \mathbb{B}_n(t)| > K\sqrt{(\ell \vee 1) \cdot \log\{[(\ell \vee 1)k + \ell] \vee 2\}} \text{ some } (s,t) \in \mathcal{S}_{k,\ell} \right) \\
&= P\left(|\mathbb{B}_n(t+s) - \mathbb{B}_n(t)| > K\sqrt{(\ell \vee 1) \cdot \log\{[(\ell \vee 1)k + \ell] \vee 2\}} \text{ some } (s,t) \in \mathcal{S}_{0,\ell} \right) \\
&\leq P\left(|\mathbb{B}_n(t)| > (K/2)\sqrt{(\ell \vee 1) \cdot \log\{[(\ell \vee 1)k + \ell] \vee 2\}} \text{ some } 0 \leq t \leq (\ell \vee 1) + \ell + 1 \right) \\
&\leq 4P\left(|\mathbb{B}_n((\ell \vee 1) + \ell + 1)| > (K/2)\sqrt{(\ell \vee 1) \cdot \log\{[(\ell \vee 1)k + \ell] \vee 2\}} \right) \\
&\leq 4 \cdot \frac{1}{\sqrt{2\pi}} \cdot \exp\left(-\frac{1}{2} \frac{(K/2)^2 (\ell \vee 1) \cdot \log\{[(\ell \vee 1)k + \ell] \vee 2\}}{(\ell \vee 1) + \ell + 1} \right) \\
&\leq 4 \cdot \frac{1}{\sqrt{2\pi}} \cdot \exp\left(-\frac{(K/2)^2 \log\{[(\ell \vee 1)k + \ell] \vee 2\}}{6} \right) = 4 \cdot \frac{1}{\sqrt{2\pi}} \cdot \{[(\ell \vee 1)k + \ell] \vee 2\}^{-K^2/24}.
\end{aligned}
$$

The third line follows since $\mathbb{B}_n(t)$ has the same distribution as $\mathbb{B}_n(t + (\ell \vee 1)k)$. The fourth line follows since, if $|\mathbb{B}_n(t+s) - \mathbb{B}_n(t)| > C$ for some C and $(s,t) \in \mathcal{S}_{0,\ell}$, we must have $|\mathbb{B}_n(t)| > C/2$ for some $0 \leq t \leq (\ell \vee 1) + \ell + 1$. The fifth line follows from the reflection principle for the Brownian motion (see Theorem 2.21 in Mörters and Peres, 2010). The sixth line uses the fact that $P(Z \geq x) \leq \frac{1}{\sqrt{2\pi}} \exp(-x^2/2)$ for $x \geq 1$ and $Z \sim N(0,1)$.

Thus,

$$P\left(\left|\mathbb{B}_n(t') - \mathbb{B}_n(t)\right| > K\sqrt{(|t-t'|\vee 1)\cdot \log(t\vee t'\vee 1)} \text{ some } 0\le t,t' < \infty\right)$$
$$\le \sum_{k=0}^{\infty}\sum_{\ell=0}^{\infty} 4\cdot\frac{1}{\sqrt{2\pi}}\cdot\{[(\ell\vee 1)k+\ell]\vee 2\}^{-K^2/24}.$$

This can be made arbitrarily small by making K large.

□

Theorem A.2 now follows since, for any constant $\varepsilon > 0$, there is a constant K such that $\sup_{\underline{h}_n\le h\le \bar{x}/A}|\hat{\mathbb{H}}_n(h) - \mathbb{H}_n(h)|$ is less than $K\{(\log n\underline{h}_n)(n\underline{h}_n)^{-1/2} + (n\underline{h}_n)^{-1/4}[\log(n\underline{h}_n)]^{1/4}\}$ with probability at least $1-\varepsilon$ asymptotically.

S1.4 Calculations for Extreme Value Limit

This section provides the calculations for the asymptotic distribution derived in Theorem A.3 in Section A.5 of the appendix.

As described in the proof of Theorem A.3, we use Theorem 12.3.5 of Leadbetter, Lindgren, and Rootzen (1983) applied to the process $\mathbb{X}(t) = \mathbb{H}(e^t)$, which is stationary, with, in the case where $k(A) \ne 0$, $\alpha = 1$ and $C = \frac{Ak(A)^2}{\int k(u)^2\,du}$ and, in the case where $k(A) = 0$, $\alpha = 2$ and $C = \frac{\int\left[k'(u)u+\frac{1}{2}k(u)\,du\right]^2 du}{2\int k(u)^2\,du}$.

In the notation of that theorem, we have

$$r(t) = cov\left(\mathbb{X}(s), \mathbb{X}(s+t)\right) = \frac{e^{\frac{1}{2}t}\int k(ue^t)k(u)\,du}{\int k(u)^2\,du}.$$

Since $r(t)$ is bounded by a constant times $e^{\frac{1}{2}t}\cdot e^{-t}$, the condition $r(t)\log t \overset{t\to\infty}{\to} 0$ holds, so it remains to verify that $r(t) = 1 - C|t|^\alpha + o(|t|^\alpha)$ with α and C given above.

Since $k(ue^t)k(u)$ has a continuous derivative with respect to t on its support, which is $[-Ae^{-t}, Ae^{-t}]$ for $t\ge 0$, it follows by Leibniz's rule that, $\frac{d}{dt}\int k(ue^t)k(u)\,du = -2Ae^{-t}k(A)k(Ae^{-t}) + \int k'(ue^t)k(u)ue^t\,du$

for $t \geq 0$, (here, we also use symmetry of k). Thus, for $t \geq 0$,

$$\frac{d}{dt_+}r(t) = \frac{e^{\frac{1}{2}t}\frac{d}{dt_+}\int k(ue^t)k(u)\,du + \frac{1}{2}e^{\frac{1}{2}t}\int k(ue^t)k(u)\,du}{\int k(u)^2\,du}$$
$$= \frac{e^{\frac{1}{2}t}\left[-2Ae^{-t}k(A)k(Ae^{-t}) + \int k'(ue^t)k(u)ue^t\,du\right] + \frac{1}{2}e^{\frac{1}{2}t}\int k(ue^t)k(u)\,du}{\int k(u)^2\,du}.$$

Thus,

$$\left.\frac{d}{dt_+}r(t)\right|_{t=0} = \frac{-2Ak(A)^2 + \int k'(u)k(u)u\,du + \frac{1}{2}\int k(u)^2\,du}{\int k(u)^2\,du} = \frac{-Ak(A)^2}{\int k(u)^2\,du}$$

where the last step follows by noting that, applying integration by parts with $k(u)u$ playing the part of u and $k'(u)du$ playing the part of dv,

$$\int k(u)k'(u)u\,du = \left[k(u)^2u\right]_{-A}^{A} - \int k(u)[k(u) + k'(u)u]\,du$$
$$= 2k(A)^2A - \int k(u)^2\,du - \int k(u)k'(u)u\,du$$

so that $\int k(u)k'(u)u\,du = k(A)^2A - \frac{1}{2}\int k(u)^2\,du$. For the case where $k(A) \neq 0$, it follows from this and a symmetric argument for $t \leq 0$ that $r(t) = 1 - C|t| - o(|t|)$ for $C = \frac{Ak(A)^2}{\int k(u)^2\,du}$ as required.

For the case where $k(A) = 0$, applying Leibniz's rule as above shows that $r(t)$ is differentiable with,

$$r'(t) = e^{\frac{1}{2}t}\frac{\int k'(ue^t)k(u)ue^t\,du + \frac{1}{2}\int k(ue^t)k(u)\,du}{\int k(u)^2\,du}.$$

Thus, $r'(0) = 0$ (using the integration by parts identity above) and $r(t)$ is twice differentiable with

$$r''(t) = e^{\frac{1}{2}t}\frac{\frac{d}{dt}\int k'(ue^t)k(u)ue^t\,du + \frac{1}{2}\frac{d}{dt}\int k(ue^t)k(u)\,du}{\int k(u)^2\,du} + \frac{1}{2}e^{\frac{1}{2}t}\frac{\int k'(ue^t)k(u)ue^t\,du + \frac{1}{2}\int k(ue^t)k(u)\,du}{\int k(u)^2\,du}.$$

We have

$$\frac{d}{dt}\int k'(ue^t)k(u)ue^t\,du = \frac{d}{dt}\int k'(v)k(ve^{-t})ve^{-t}\,dv$$
$$= \int k'(v)k'(ve^{-t})(-ve^{-t})ve^{-t}\,dv - \int k'(v)k(ve^{-t})ve^{-t}\,dv$$

and $\frac{d}{dt}\int k(ue^t)k(u)\,du = \int k'(ue^t)k(u)ue^t\,du$, so this gives

$$r''(t) = e^{\frac{1}{2}t}\frac{-\int k'(v)k'(ue^{-t})u^2e^{-2t}\,du - \frac{1}{2}\int k'(ue^t)k(u)ue^t\,du}{\int k(u)^2\,du} + \frac{1}{2}e^{\frac{1}{2}t}\frac{\int k'(ue^t)k(u)ue^t\,du + \frac{1}{2}\int k(ue^t)k(u)\,du}{\int k(u)^2\,du}.$$

Thus,

$$r''(0) = \frac{-\int [k'(u)u]^2\,du + \frac{1}{4}\int k(u)^2\,du}{\int k(u)^2\,du}.$$

Since, by the integration by parts argument above, $\frac{1}{4}\int k(u)^2\,du = \frac{1}{2}\int k(u)^2\,du - \frac{1}{4}\int k(u)^2\,du = -\int k(u)k'(u)u\,du - \frac{1}{4}\int k(u)^2\,du$, this is equal to

$$\frac{-\int [k'(u)u]^2\,du - \int k(u)k'(u)u\,du - \frac{1}{4}\int k(u)^2\,du}{\int k(u)^2\,du} = -\frac{\int \left[k'(u)u + \frac{1}{2}k(u)\right]^2\,du}{\int k(u)^2\,du}$$

which gives the required expansion with C given by one half of the negative of the above display and $\alpha = 2$.

S1.5 Delta Method

We state some results that allow us to obtain influence function representations with the necessary uniform rate for differentiable functions of estimators. These results amount to applying the delta method to our setting and keeping track of the uniform rates.

Let $\hat{\beta}(h)$ be an estimator of a parameter $\beta(h) \in \mathbb{R}^{d_\beta}$ with influence function representation

$$\sqrt{nh}(\hat{\beta}(h) - \beta(h)) = \frac{1}{\sqrt{nh}}\sum_{i=1}^{n}\psi_\beta(W_i, h)k(X_i/h) + R_{1,n}(h)$$

for some function ψ_β and a kernel function k, where $\psi_\beta(W_i, h)k(X_i/h)$ has mean zero and $\sup_{\underline{h}_n \le h \le \overline{h}} |R_{1,n}(h)| = o_P(1/\sqrt{\log\log \underline{h}_n^{-1}})$. Let g be a function from $\mathbb{R}^{d_\beta}$ to $\mathbb{R}^{d_\theta}$ and consider the parameter $\theta(h) = g(\beta(h))$ and the estimator $\hat{\theta}(h) = g(\hat{\beta}(h))$.

Let $\hat{V}_\beta(h)$ be an estimate of $V_\beta(h) = \frac{1}{h}E\psi_\beta(W_i, h)\psi_\beta(W_i, h)'k(X_i/h)^2$, the (pointwise in h)

asymptotic variance of $\hat{\beta}(h)$. A natural estimator of the asymptotic variance $V_\theta(h)$ of $\hat{\theta}$ is

$$\hat{V}_\theta(h) = D_g(\hat{\beta}(h))'\hat{V}_\beta(h)D_g(\hat{\beta}(h))'.$$

Lemma S1.11. *Suppose that $\beta(h)$ is bounded uniformly over $h \le \overline{h}_n$ where $\overline{h}_n = \mathcal{O}(1)$ and*

(i) For large enough n, g is differentiable on an open set containing the range of $\beta(h)$ over $h \le \overline{h}_n$, with Lipschitz continuous derivative D_g.

(ii) ψ_β and k are bounded, k has finite support, and the class of functions $(w, x) \mapsto \psi_\beta(w, h)k(x/h)$ has polynomial uniform covering number.

(iii) $|X_i|$ has a bounded density on $[0, \overline{h}_n]$ for large enough n.

Then, if $n\underline{h}_n/(\log\log n)^3 \to \infty$,

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \left| \sqrt{nh}(\hat{\theta}(h) - \theta(h)) - \frac{1}{\sqrt{nh}} \sum_{i=1}^n D_g(\beta(h))\psi_\beta(W_i, h)k(X_i/h) \right| = o_P\left(1 \Big/ \sqrt{\log\log \underline{h}_n^{-1}}\right).$$

If, in addition, $\sup_{\underline{h}_n \le h \le \overline{h}_n} \|\hat{V}_\beta(h) - V_\beta(h)\| \xrightarrow{p} 0$, then, for some constant K and some $R_{n,2}(h)$ with $\sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{\sqrt{nh}}{\sqrt{\log\log h^{-1}}} |R_{n,2}(h)| = \mathcal{O}_P(1)$,

$$\left\|\hat{V}_\theta(h) - V_\theta(h)\right\| \le K \left\|\hat{V}_\beta(h) - V_\beta(h)\right\| + R_{n,2}(h)$$

for all $\underline{h}_n \le h \le \overline{h}_n$ with probability approaching one.

Proof. By a first order Taylor expansion, we have, for some $\beta^*(h)$ with $\|\beta^*(h) - \beta(h)\| \le \|\hat{\beta}(h) - \beta(h)\|$,

$$\begin{aligned}
&\sqrt{nh}(\hat{\theta}(h) - \theta(h)) = \sqrt{nh}(g(\hat{\beta}(h)) - g(\beta(h))) = \sqrt{nh}D_g(\beta^*(h))(\hat{\beta}(h) - \beta(h)) \\
&= D_g(\beta^*(h)) \frac{1}{\sqrt{nh}} \sum_{i=1}^n \psi_\beta(W_i, h)k(X_i/h) + D_g(\beta^*(h))R_{1,n}(h) \\
&= \frac{1}{\sqrt{nh}} \sum_{i=1}^n D_g(\beta(h))\psi_\beta(W_i, h)k(X_i/h) + [D_g(\beta^*(h)) - D_g(\beta(h))] \frac{1}{\sqrt{nh}} \sum_{i=1}^n \psi_\beta(W_i, h)k(X_i/h) \\
&+ D_g(\beta^*(h))R_{1,n}(h)
\end{aligned}$$

Applying Lemma A.2, $\hat{\beta}(h) - \beta(h)$ is $\mathcal{O}_P(\sqrt{\log\log h^{-1}}/\sqrt{nh})$ uniformly over $\underline{h}_n \leq h \leq \overline{h}_n$ and $\frac{1}{\sqrt{nh}}\sum_{i=1}^n \psi_\beta(W_i,h)k(X_i/h)$ is $\mathcal{O}_P(\sqrt{\log\log h^{-1}})$ uniformly over $\underline{h}_n \leq h \leq \overline{h}_n$. so that, by the Lipschitz condition on D_g, the second term is $\mathcal{O}_P(\log\log h^{-1}/\sqrt{nh})$ uniformly over $\underline{h}_n \leq h \leq \overline{h}_n$, which is $o_P(1/\sqrt{\log\log \underline{h}_n^{-1}})$ uniformly over $\underline{h}_n \leq h \leq \overline{h}_n$ since $\sqrt{n\underline{h}_n}/(\log\log \underline{h}_n)^{3/2} \to \infty$. The last term is $o_P(1/\sqrt{\log\log h_n^{-1}})$ uniformly over $\underline{h}_n \leq h \leq \overline{h}_n$ by the conditions on $R_{1,n}(h)$, the uniform consistency of $\hat{\beta}(h)$ and the Lipschitz condition on D_g.

For the second claim, note that

$$
\begin{aligned}
&\hat{V}_\theta - V_\theta = D_g(\hat{\beta}(h))\hat{V}_\beta(h)D_g(\hat{\beta}(h))' - D_g(\beta(h))V_\beta(h)D_g(\beta(h))' \\
&= [D_g(\hat{\beta}(h)) - D_g(\beta(h))]\hat{V}_\beta(h)D_g(\hat{\beta}(h))' + D_g(\beta(h))[\hat{V}_\beta(h) - V_\beta(h)]D_g(\hat{\beta}(h))' \\
&+ D_g(\beta(h))V_\beta(h)[D_g(\hat{\beta}(h)) - D_g(\beta(h))]'.
\end{aligned}
$$

The first and last terms converge at a $\sqrt{\log\log h^{-1}}/\sqrt{nh}$ rate uniformly over $\underline{h}_n \leq h \leq \overline{h}_n$ by Lemma A.2 and the Lipschitz continuity on D_g. The second term is bounded by a constant times $\|\hat{V}_\beta(h) - \hat{V}_\beta(h)\|$ uniformly over $\underline{h}_n \leq h \leq \overline{h}$ with probability approaching one by the uniform consistency of $\hat{\beta}(h)$ and the Lipschitz continuity of D_g.

□

S1.6 Sufficient Conditions Based on Non-normalized Influence Function

In some cases, it will be easier to verify the conditions for an influence function approximation to $\sqrt{nh}(\hat{\theta}(h) - \theta(h))$ rather than the normalized version $\sqrt{nh}(\hat{\theta}(h) - \theta(h))/\hat{\sigma}(h)$. The following lemma is useful in these cases.

Lemma S1.12. *Suppose that the following conditions hold for some $\tilde{\psi}(W_i,h)$.*

(i) $E\tilde{\psi}(W_i,h)k(X_i/h) = 0$ and k is bounded and symmetric with finite support $[-A,A]$.

(ii) $|X_i|$ has a density $f_{|X|}$ with $f_{|X|}(0) > 0$, $\tilde{\psi}(W_i,h)k(X_i/h)$ is bounded uniformly over $h \leq \underline{h}_n$ and, for some deterministic function $\ell(h)$ with $\ell(h)\log\log h^{-1} \to 0$ as $h \to 0$, the following expressions are bounded by $\ell(t)$: $|f_{|X|}(t) - f_{|X|}(0)|$, $|E[\tilde{\psi}(W_i,0)||X_i| = t] - E[\tilde{\psi}(W_i,0)||X_i| = 0]|$, $|var[\tilde{\psi}(W_i,0)||X_i| = t] - var[\tilde{\psi}(W_i,0)||X_i| = 0]|$ and $|(\tilde{\psi}(W_i,t) - \tilde{\psi}(W_i,0))k(X_i/h)|$.

Let $\sigma^2(h) = \frac{1}{h}var(\tilde{\psi}(W_i,h)k(X_i/h))$ *for* $h > 0$ *and* $\sigma^2(0) = var\left[\tilde{\psi}(W_i,0)||X_i| = 0\right] f_{|X|}(0) \int_{u=0}^{\infty} k(u)^2\, du$. *Let* $\psi(W_i,h) = \tilde{\psi}(W_i,h)/\sigma(h)$ *so that* $\frac{1}{h}var[\psi(W_i,h)k(X_i/h)] = 1$. *Suppose that* $var\left[\tilde{\psi}(W_i,0)||X_i| = 0\right] > 0$. *Then the above assumptions hold with* $\tilde{\psi}$ *replaced by* ψ *for* h *small enough and with* $\ell(t)$ *possibly redefined.*

Proof. First, note that the only condition we need to verify is the one involving $|[\psi(W_i,h) - \psi(W_i,0)]k(X_i/h)|$, since the remaining conditions are only changed by multiplication by a constant when $\tilde{\psi}$ is replaced by ψ. Note that

$$\sigma^2(h) - \frac{1}{h}var(\tilde{\psi}(W_i,0)k(X_i/h)) = \frac{1}{h}var(\tilde{\psi}(W_i,h)k(X_i/h)) - \frac{1}{h}var(\tilde{\psi}(W_i,0)k(X_i/h))$$
$$= \frac{1}{h}var\{[\tilde{\psi}(W_i,h) - \tilde{\psi}(W_i,0)]k(X_i/h)\} + 2\frac{1}{h}cov\{[\tilde{\psi}(W_i,h) - \tilde{\psi}(W_i,0)]k(X_i/h), \tilde{\psi}(W_i,0)k(X_i/h)\}.$$

Since $|(\tilde{\psi}(W_i,h) - \tilde{\psi}(W_i,0))k(X_i/h)| \leq \ell(h)I(|X_i| \leq Ah)$ and $\tilde{\psi}(W_i,h)k(X_i/h)$ and $\tilde{\psi}(W_i,0)k(X_i/h)$ are bounded, the last two terms are bounded by a constant times $\ell(h)\frac{1}{h}EI(|X_i| \leq Ah)$, which is bounded by a constant times $\ell(h)$ by the assumption on the density of $|X_i|$.

Thus, let us consider

$$\frac{1}{h}var(\tilde{\psi}(W_i,0)k(X_i/h))$$
$$= \frac{1}{h}\int_{x=0}^{\infty} var\left[\tilde{\psi}(W_i,0)||X_i| = x\right] k(x/h)^2 f_{|X|}(x)\, dx + \frac{1}{h}var\left\{E\left[\tilde{\psi}(W_i,0)||X_i|\right] k(X_i/h)\right\}.$$

Arguing as in the proof of Lemma A.6 (using the fact that $E\tilde{\psi}(W_i,h)k(X_i/h) = 0$ and taking limits), it can be seen that $E\left[\tilde{\psi}(W_i,0)||X_i| = 0\right] = 0$ under these conditions. Thus, the last term is bounded by $\ell(Ah)^2\frac{1}{h}Ek(X_i/h)^2$. The first term is equal to $var\left[\tilde{\psi}(W_i,0)||X_i| = 0\right] f_{|X|}(0) \int_{u=0}^{\infty} k(u)^2\, du$ plus a term that is bounded by a constant times $\ell(Ah)$.

It follows that, letting $\sigma^2(0) = var\left[\tilde{\psi}(W_i,0)||X_i| = 0\right] f_{|X|}(0) \int_{u=0}^{\infty} k(u)^2\, du$ as defined above, we have, for some constant K, $|\sigma^2(h) - \sigma^2(0)| \leq K\ell(Ah)$. Thus,

$$|[\psi(W_i,h) - \psi(W_i,0)]k(X_i/h)|$$
$$\leq \frac{1}{\sigma(0)}\left|[\tilde{\psi}(W_i,h) - \tilde{\psi}(W_i,0)]k(X_i/h)\right| + |\tilde{\psi}(W_i,h)k(X_i/h)| \cdot \left|\frac{1}{\sigma(h)} - \frac{1}{\sigma(0)}\right|.$$

The first term is bounded by a constant times $\ell(h)$ by assumption. The last term is bounded by a constant times $|\sigma^2(h) - \sigma^2(0)|$, which is bounded by a constant times $\ell(Ah)$ as shown above. $\square$

S2 Local polynomial estimators: regression discontinuity/estimation at the boundary

This section gives primitive conditions for smooth functions of estimates based on local polynomial estimates at the boundary, or at a discontinuity in the regression function. The results are used in Section S3 below to verify the conditions of Theorem 3.1 for the applications in Section 5 in the main text. Throughout this section, we consider a setup with $\{(X_i, Y_i')'\}_{i=1}^n$ iid with X_i a real valued random variable and Y_i taking values in $\mathbb{R}^{d_Y}$. We consider smooth functions of the left and right hand limits of the regression function at a point, which we normalize to be zero.

Let $(\hat{\beta}_{u,j,1}(h), \hat{\beta}_{u,j,2}(h)/h, \ldots, \hat{\beta}_{u,j,r+1}(h)/h^r)$ be the coefficients of an rth order local polynomial estimate of $E[Y_{i,j}|X_i = 0_+]$ based on the subsample with $X_i \geq 0$ with a kernel function k^*. Similarly, let $(\hat{\beta}_{\ell,j,1}(h), \hat{\beta}_{\ell,j,2}(h)/h, \ldots, \hat{\beta}_{\ell,j,r+1}(h)/h^r)$ be the coefficients of an rth order local polynomial estimate of $E[Y_{i,j}|X_i = 0_-]$ based on the subsample with $X_i < 0$, where the polynomial is taken in $|X_i|$ rather than X_i (this amounts to multiplying even elements of $\beta_{\ell,j}$ by -1). The scaling by powers of h is used to handle the different rates of convergence of the different coefficients. Let $p(x) = (1, x, x^2, \ldots, x^r)'$, and define $\hat{\beta}_{u,j} = (\hat{\beta}_{u,j,1}(h), \hat{\beta}_{u,j,2}(h), \ldots, \hat{\beta}_{u,j,r+1}(h))$ and $\hat{\beta}_{\ell,j} = (\hat{\beta}_{\ell,j,1}(h), \hat{\beta}_{\ell,j,2}(h), \ldots, \hat{\beta}_{\ell,j,r+1}(h))$. Let $p(x) = (1, x, x^2, \ldots, x^r)'$. Then $\hat{\beta}_{u,j}$ minimizes

$$\sum_{i=1}^{n} (Y_{i,j} - p(|X_i/h|)'\beta_{u,j})^2 I(X_i \geq 0) k^*(X_i/h)$$

and $\hat{\beta}_{\ell,j}$ minimizes

$$\sum_{i=1}^{n} (Y_{i,j} - p(|X_i/h|)'\beta_{u,j})^2 I(X_i < 0) k^*(X_i/h).$$

Define

$$\Gamma_u(h) = \tfrac{1}{h} Ep(|X_i/h|)p(|X_i/h|)'k^*(X_i/h)I(X_i \geq 0),$$
$$\Gamma_\ell(h) = \tfrac{1}{h} Ep(|X_i/h|)p(|X_i/h|)'k^*(X_i/h)I(X_i < 0),$$
$$\hat{\Gamma}_u(h) = \tfrac{1}{nh} \textstyle\sum_{i=1}^n p(|X_i/h|)p(|X_i/h|)'k^*(X_i/h)I(X_i \geq 0) \qquad \text{and}$$
$$\hat{\Gamma}_\ell(h) = \tfrac{1}{nh} \textstyle\sum_{i=1}^n p(|X_i/h|)p(|X_i/h|)'k^*(X_i/h)I(X_i < 0).$$

Let $\mu_{k^*,\ell} = \int_0^\infty u^\ell k^*(u)\, du$, and let M be the matrix with i,jth element given by $\mu_{k^*,i+j-2}$.

Let $\hat{\alpha}_u(h) = (\hat{\beta}_{1,1,u}(h), \dots \hat{\beta}_{1,d_Y,u}(h))'$ and $\hat{\alpha}_\ell(h) = (\hat{\beta}_{1,1,\ell}(h), \dots \hat{\beta}_{1,d_Y,\ell}(h))'$, and similarly for $\alpha_u(h)$ and $\alpha_\ell(h)$ (i.e. α_u and α_ℓ contain the constant terms in the local polynomial regressions for each j). Let $\hat{\alpha}(h) = (\hat{\alpha}_u(h)', \hat{\alpha}_\ell(h)')$ and $\alpha(h) = (\alpha_u(h)', \alpha_\ell(h)')$. We are interested in $\theta(h) = g(\alpha(h))$ for a differentiable function g from $\mathbb{R}^{2d_Y}$ to $\mathbb{R}$, and an estimator $\hat{\theta}(h) = \hat{g}(\alpha(h))$. We consider standard errors defined by the delta method applied to the robust covariance matrix formula obtained by treating the local linear regressions as a system of $2d_Y$ weighted least squares regressions. Let $\nu_u(h) = e_1'\Gamma_u(h)^{-1}$ and let $\nu_\ell(h) = e_1'\Gamma_\ell(h)^{-1}$. Let $\hat{\nu}_u(h) = e_1'\hat{\Gamma}_u(h)^{-1}$ and let $\nu_\ell(h) = e_1'\hat{\Gamma}_\ell(h)^{-1}$. Let $\psi_\alpha(X_i, Y_i, h)$ be the $(2d_Y) \times 1$ random vector with jth element given by

$$\psi_{\alpha,j}(X_i, Y_i, h) = \begin{cases} \nu_u(h)p(|X_i/h|)[Y_{i,j} - p(|X_i/h|)'\beta_{u,j}(h)]I(X_i \geq 0) & \text{if } j = 1, \dots, d_Y, \\ \nu_\ell(h)p(|X_i/h|)[Y_{i,j-d_Y} - p(|X_i/h|)'\beta_{\ell,j-d_Y}(h)]I(X_i < 0) & \text{if } j = d_Y + 1, \dots, 2d_Y. \end{cases}$$

Let $\hat{\psi}_\alpha(X_i, Y_i, h)$ be defined analogously,

$$\hat{\psi}_{\alpha,j}(X_i, Y_i, h) = \begin{cases} \hat{\nu}_u(h)p(|X_i/h|)[Y_{i,j} - p(|X_i/h|)'\hat{\beta}_{u,j}(h)]I(X_i \geq 0) & \text{if } j = 1, \dots, d_Y, \\ \hat{\nu}_\ell(h)p(|X_i/h|)[Y_{i,j-d_Y} - p(|X_i/h|)'\hat{\beta}_{\ell,j-d_Y}(h)]I(X_i < 0) & \text{if } j = d_Y + 1, \dots, 2d_Y. \end{cases}$$

Let

$$V_\alpha(h) = \frac{1}{h} E\psi_\alpha(X_i, Y_i, h)\psi_\alpha(X_i, Y_i, h)'k^*(X_i/h)^2$$

and let

$$\hat{V}_\alpha(h) = \frac{1}{h} E_n \hat{\psi}_\alpha(X_i, Y_i, h)\hat{\psi}_\alpha(X_i, Y_i, h)' k^*(X_i/h)^2.$$

Let $\hat{\sigma}(h) = D_g(\hat{\alpha}(h))\hat{V}_\alpha(h)D_g(\hat{\alpha}(h))'$, and $\sigma(h) = D_g(\alpha(h))V_\alpha(h)D_g(\alpha(h))'$, where D_g is the derivative of g.

We make the following assumption throughout this section. In the following assumption, $\ell(t)$ is an arbitrary nondecreasing function satisfying $\lim_{t\downarrow 0} \ell(t) \log\log t^{-1} = 0$.

Assumption S2.1. *(i) X_i has a density $f_X(x)$ with $|f_X(x) - f_{X,-}| \leq \ell(x)$ for $x < 0$ and $|f_X(x) - f_{X,+}| \leq \ell(x)$ for some $f_{X,+} > 0$ and $f_{X,-} > 0$.*

(ii) Y_i is bounded and, for some matrices Σ_- and Σ_+ and vectors $\tilde{\mu}_-$ and $\tilde{\mu}_+$, $\tilde{\Sigma}(x) = var(Y_i|X_i = x)$ and $\tilde{\mu}(x) = E(Y_i|X_i = x)$ satisfy $\|\tilde{\Sigma}(x) - \Sigma_+\| \leq \ell(x)$ and $\|\tilde{\mu}(x) - \tilde{\mu}_+\| \leq \ell(x)$ for $x > 0$ and $\|\tilde{\Sigma}(x) - \Sigma_-\| \leq \ell(x)$ and $\|\tilde{\mu}(x) - \tilde{\mu}_-\| \leq \ell(x)$ for $x < 0$.

(iii) k^* is symmetric with finite support $[-A, A]$, is bounded with a bounded, uniformly continuous first derivative on $(0, A)$, and satisfies $\int k(u)\, du \neq 0$, and the matrix M is invertible.

(iv) D_g is bounded and is Lipschitz continuous on an open set containing the range of $\alpha(h)$ over $\overline{h}_n$ for n large enough.

(v) $D_{g,u}(\alpha(0))\tilde{\Sigma}_+ D_{g,u}(\alpha(0)) > 0$ or $D_{g,\ell}(\alpha(0))\tilde{\Sigma}_- D_{g,u}(\ell) > 0$.

(vi) $\overline{h}_n = \mathcal{O}(1)$ and $n\underline{h}_n/(\log\log n)^3 \to \infty$.

We prove the following theorem.

Theorem S2.1. *Under Assumption S2.1, Assumptions 3.1 and Assumption 3.2 hold with $k(u) = e_1' M^{-1} p(|u|) k^*(u)$ and ψ defined below so long as $n\underline{h}_n/(\log\log \underline{h}_n^{-1})^3 \to \infty$ and $\overline{h}_n$ is small enough for large n.*

Throughout the following, we assume that $\overline{h}_n$ is small enough so that $\|\Gamma_u(h)^{-1}\|$ and $\|\Gamma_\ell(h)^{-1}\|$ are bounded uniformly over $h \leq \overline{h}_n$ for large enough n (this will hold for small enough $\overline{h}_n$ by Lemma S2.4 below).

Lemma S2.1. *Suppose that Assumption S2.1 holds. Then*

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{\sqrt{nh}}{\sqrt{\log\log h^{-1}}} \left\| \hat{\Gamma}_u(h) - \Gamma_u(h) \right\| = \mathcal{O}_P(1),$$

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{\sqrt{nh}}{\sqrt{\log\log h^{-1}}} \left\| \hat{\Gamma}_u(h)^{-1} - \Gamma_u(h)^{-1} \right\| = \mathcal{O}_P(1),$$

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{nh}{\log\log h^{-1}} \Big\| \hat{\beta}_{u,j}(h) - \beta_{u,j}(h) \\ - \frac{1}{h} E_n \Gamma_u(h)^{-1} p(X_i/h) k^*(X_i/h)[Y_i - p(X_i/h)'\beta(h)] I(X_i \ge 0) \Big\| = \mathcal{O}_P(1),$$

and

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{\sqrt{nh}}{\sqrt{\log\log h^{-1}}} \left\| \hat{\beta}_{u,j}(h) - \beta_{u,j}(h) \right\| = \mathcal{O}_P(1)$$

for each j. The same holds with $I(X_i \ge 0)$ replaced by $I(X_i < 0)$, Γ_u replaced by Γ_ℓ, $\hat{\Gamma}_u$ replaced by $\hat{\Gamma}_\ell$, etc.

Proof. The first display follows from Lemma A.2. For the second display, note that $\hat{\Gamma}(h)^{-1} - \Gamma(h)^{-1} = -\hat{\Gamma}(h)^{-1}(\hat{\Gamma}(h) - \Gamma(h))\Gamma(h)^{-1}$, so $\|\hat{\Gamma}(h)^{-1} - \Gamma(h)^{-1}\| \le \|\hat{\Gamma}(h)^{-1}\| \|\hat{\Gamma}(h) - \Gamma(h)\| \|\Gamma(h)^{-1}\|$. $\|\Gamma(h)^{-1}\|$ is bounded by assumption and $\|\hat{\Gamma}(h)^{-1}\|$ is $\mathcal{O}_P(1)$ uniformly over $\underline{h}_n \le h \le \overline{h}_n$ by this and the first display in the lemma. For the third display, note that

$$\hat{\beta}_{u,j}(h) - \beta_{u,j}(h) = \hat{\Gamma}_u(h)^{-1} \frac{1}{h} E_n p(X_i/h) k^*(X_i/h)[Y_i - p(X_i/h)'\beta(h)] I(X_i \ge 0).$$

Thus,

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{nh}{\log\log h^{-1}} \left\| \hat{\beta}_{u,j}(h) - \beta_{u,j}(h) - \frac{1}{h} E_n \Gamma_u(h)^{-1} p(X_i/h) k^*(X_i/h)[Y_i - p(X_i/h)'\beta(h)] I(X_i \ge 0) \right\|$$
$$\le \sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{\sqrt{nh}}{\sqrt{\log\log h^{-1}}} \left\| \hat{\Gamma}_u(h)^{-1} - \Gamma_u(h)^{-1} \right\|$$
$$\cdot \sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{\sqrt{nh}}{\sqrt{\log\log h^{-1}}} \left\| \frac{1}{h} E_n p(X_i/h) k^*(X_i/h)[Y_i - p(X_i/h)'\beta(h)] I(X_i \ge 0) \right\|.$$

The first term is $\mathcal{O}_P(1)$ by the second display in the lemma. The second term is $\mathcal{O}_P(1)$ by Lemma A.2. The last display in the lemma follows from the third display and Lemma A.2. □

Applying the above lemma, we obtain the following.

Lemma S2.2. *Under Assumption S2.1,*

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{nh}{\log\log h^{-1}} \left\| \hat{\alpha}(h) - \alpha(h) - \frac{1}{nh} \sum_{i=1}^{n} \psi_\alpha(X_i, Y_i, h) k^*(X_i/h) \right\| = \mathcal{O}_P(1)$$

and

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{\sqrt{nh}}{\sqrt{\log\log h^{-1}}} \left\| \hat{V}_\alpha(h) - V_\alpha(h) \right\| = \mathcal{O}_P(1).$$

Proof. The first claim follows by Lemma S2.1. The second claim follows by using the fact that $\hat{V}_\alpha(h)$ is a Lipschitz continuous function of the $\hat{\beta}$ and $\hat{\nu}$ terms and terms that can be handled with Lemma A.2. □

Lemma S2.3. *Suppose that Assumption S2.1 holds. Then*

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \sqrt{nh} \left\| \hat{\theta}(h) - \theta(h) - \frac{1}{\sqrt{nh}} \sum_{i=1}^{n} D_g(\alpha(h)) \psi_\alpha(X_i, Y_i, h) k^*(X_i/h) \right\| = o_P\left(1/\sqrt{\log\log \underline{h}_n^{-1}}\right)$$

and

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \frac{\sqrt{nh}}{\sqrt{\log\log h^{-1}}} \left\| \hat{\sigma}(h) - \sigma(h) \right\| = \mathcal{O}_P(1).$$

Proof. By Lemma S2.2,

$$\sup_{\underline{h}_n \le h \le \overline{h}_n} \left\| \sqrt{nh}\,(\hat{\alpha}(h) - \alpha(h)) - \frac{1}{\sqrt{nh}} \sum_{i=1}^{n} \psi_\alpha(X_i, Y_i, h) k^*(X_i/h) \right\|$$

$$= \mathcal{O}_P\left(\sup_{\underline{h} \le h \le \overline{h}_n} (\log\log h^{-1})/\sqrt{nh} \right) = \mathcal{O}_P\left((\log\log \underline{h}_n^{-1})/\sqrt{n\underline{h}_n} \right) = o_P\left(1/\sqrt{\log\log \underline{h}_n^{-1}} \right)$$

since $(\log\log \underline{h}_n^{-1})^{3/2}/\sqrt{n\underline{h}_n} \to 0$. Thus, the result follows by Lemma S1.11.

□

Let $m_j(x,h) = p(x/h)'\beta_{u,j}(h)$ for $x \ge 0$ and $m_j(x,h) = p(x/h)'\beta_{\ell,j-d_Y}(h)$ for $x < 0$. Let $D_{g,u}(\alpha)$ be the row vector with the first d_Y elements of $D_g(\alpha)$, and let $D_{g,\ell}(\alpha)$ be the row vector with the remaining d_Y elements. With this notation, we have

$$D_g(\alpha(h))\psi_\alpha(X_i, Y_i, h)$$
$$= \left\{ I(X_i \ge 0)\nu_u(h) p(|X_i/h|) D_{g,u}(\alpha(h)) + I(X_i < 0)\nu_\ell(h) p(|X_i/h|) D_{g,\ell}(\alpha(h)) \right\} [Y_i - m(X_i, h)].$$

Let $\gamma_{u,j}(h) = \frac{1}{h} E Y_{i,j} p(|X_i/h|) k^*(X_i/h) I(X_i \ge 0)$ and $\gamma_{u,\ell}(h) = \frac{1}{h} E Y_{i,j} p(|X_i/h|) k^*(X_i/h) I(X_i < 0)$. Let $\gamma_{u,j}(0)$ be the $(r+1) \times 1$ vector with qth element given by $f_{X,+}\tilde{\mu}_{+,j}\mu_{k^*,q}$ and let Let $\gamma_{\ell,j}(0)$ be the $(r+1) \times 1$ vector with qth element given by $f_{X,-}\tilde{\mu}_{-,j}\mu_{k^*,q}$. Let $\alpha(0) = (\tilde{\mu}_+', \tilde{\mu}_-')'$ (it will be shown below that $\lim_{h\to 0}\alpha(h) = \alpha(0)$).

We now verify the conditions of the main result with $k(u) = e_1' M^{-1} p(|u|) k^*(u)$ and

$$\psi(W_i, h) = \frac{D_g(\alpha(h))\psi_\alpha(X_i, Y_i, h)}{e_1' M^{-1} p(|X_i/h|)\sigma(h)}$$

for $h > 0$ and

$$\psi(W_i, 0) = \frac{1}{\sigma(0)} \left[D_{g,u}(\alpha(0)) f_{X,+}^{-1}(Y_i - \mu_+) I(X_i \ge 0) + D_{g,\ell}(\alpha(0)) f_{X,-}^{-1}(Y_i - \mu_-) I(X_i < 0) \right]$$

where $\sigma^2(0) = \lim_{h\to 0}\sigma^2(h)$ (this choice of $\psi(W_i, 0)$ will be justified by the calculations below).

Lemma S2.4. *Under Assumption S2.1, for some constant K,*

$$\|\Gamma_u(h) - f_{X,+}M\| \leq K\ell(Ah),$$
$$\|\Gamma_\ell(h) - f_{X,-}M\| \leq K\ell(Ah),$$
$$\|\gamma_u(h) - \gamma_u(0)\| \leq K\ell(Ah),$$
$$\text{and} \qquad \|\gamma_\ell(h) - \gamma_\ell(0)\| \leq K\ell(Ah).$$

Proof. We have

$$\gamma_{u,j}(h) = \frac{1}{h}EY_{i,j}p(|X_i/h|)k^*(X_i/h)I(X_i \geq 0) = \frac{1}{h}\int_{x=0}^{\infty} \tilde{\mu}_j(x)p(x/h)k^*(x/h)f_X(x)\,dx$$
$$= \int_{x=0}^{\infty} \tilde{\mu}_j(uh)p(u)k^*(u)f_X(uh)\,dx.$$

Thus, by boundedness of k^*, $\|\gamma_{u,j}(h) - \gamma_{u,j}(0)\|$ is bounded by a constant times $\sup_{0\leq x\leq Ah} |\tilde{\mu}_j(x)f_X(x) - \tilde{\mu}_{+,j}f_{X,+}|$, which is bounded by a constant times $\ell(Ah)$ by assumption. Similarly,

$$\Gamma_{u,j,m}(h) = \frac{1}{h}E(X_i/h)^{j+m-2}k^*(X_i/h)I(X_i \geq 0) = \frac{1}{h}\int_{x=0}^{\infty}(x/h)^{j+m-2}k^*(x/h)f_X(x)\,dx$$
$$= \int_{x=0}^{\infty} u^{j+m-2}k^*(u)f_X(uh)\,du,$$

so $|\Gamma_{u,j,m}(h) - f_{X,+}M_{j,m}|$ is bounded by a constant times $\sup_{0\leq x\leq Ah} |f_X(x) - f_{X,+}| \leq \ell(Ah)$. The proof for Γ_ℓ and γ_ℓ is similar. □

Note that $\beta_{u,j}(h) = \Gamma_u(h)^{-1}\gamma_{u,j}(h) \to \tilde{\mu}_{+,j}M^{-1}(1, \mu_{k^*,1}, \ldots, \mu_{k^*,r})' = \tilde{\mu}_{+,j}(1, 0, \ldots, 0)'$ as $h \to 0$, where the last equality follows since $M^{-1}(1, \mu_{k^*,1}, \ldots, \mu_{k^*,r})'$ is the first column of $M^{-1}M = I_{r+1}$ (the second through rth elements of $\beta_{u,j}$ are given by the corresponding coefficients of the local polynomial scaled by powers of h, so this is a result of the fact that the coefficients of the local polynomial do not increase too quickly as $h \to 0$). By these calculations and Lemma S2.4, we obtain the following.

Lemma S2.5. *Under Assumption S2.1, for some constant K and h small enough,*

$$\left|\beta_{u,j}(h) - \tilde{\mu}_{+,j}(1, 0, \ldots, 0)'\right| \leq K\ell(Ah),$$
$$\text{and} \qquad \left|\beta_{\ell,j}(h) - \tilde{\mu}_{-,j}(1, 0, \ldots, 0)'\right| \leq K\ell(Ah).$$

Proof. The result is immediate from Lemma S2.4, the fact that $\|\Gamma_u(h)^{-1}\|$ and $\|\Gamma_\ell(h)^{-1}\|$ are bounded uniformly over small enough h (which follows from Lemma S2.4 and invertibility of M) and fact that the function that takes Γ and γ to $\Gamma^{-1}\gamma$ is Lipschitz over Γ and γ with Γ^{-1} and γ bounded. □

Note that, since $\alpha(h)$ is made up of the first component of each of the $\beta_{u,j}(h)$ and $\beta_{\ell,j}(h)$ vectors, the above lemma also implies that $|\alpha(h) - \alpha(0)| \leq K\ell(Ah)$ for $\alpha(0)$ defined above. For convenience, let us also define $\beta_{u,j}(0)$ and $\beta_{\ell,j}(0)$ to be the limits of $\beta_{u,j}(h)$ and $\beta_{\ell,j}(h)$ derived above.

Lemma S2.6. *Under Assumption S2.1, for some constant K and h small enough,*

$$\|\nu_u(h) - e_1' M^{-1} f_{X,+}^{-1}\| \leq K\ell(Ah)$$
$$\text{and} \qquad \|\nu_\ell(h) - e_1' M^{-1} f_{X,-}^{-1}\| \leq K\ell(Ah).$$

Proof. The result is immediate from Lemma S2.4 and the boundedness of $\|\Gamma_u(h)^{-1}\|$ and $\|\Gamma_\ell(h)^{-1}\|$ over small enough h. □

Lemma S2.7. *Under Assumption S2.1, for some constant K and h small enough,*

$$|[\sigma(h)\psi(W_i, h) - \sigma(0)\psi(W_i, 0)]\, k(X_i/h)| \leq K\ell(Ah).$$

Proof. We have

$$\begin{aligned}
&[\sigma(h)\psi(W_i, h) - \sigma(0)\psi(W_i, 0)]\, k(X_i/h) \\
&= D_g(\alpha(h))\psi_\alpha(X_i, Y_i, h)k^*(X_i/h) \\
&- \left[D_{g,u}(\alpha(0)) f_{X,+}^{-1}(Y_i - \mu_+)I(X_i \geq 0) + D_{g,\ell}(\alpha(0)) f_{X,-}^{-1}(Y_i - \mu_-)I(X_i < 0)\right] e_1' M^{-1} p(|X_i/h|)k^*(X_i/h) \\
&= D_g(\alpha(h))\psi_\alpha(X_i, Y_i, h)k^*(X_i/h) - D_g(\alpha(0))\tilde{\psi}_\alpha(X_i, Y_i, h)k^*(X_i/h)
\end{aligned}$$

where the first d_Y columns of $\tilde{\psi}_\alpha(X_i, Y_i, h)$ are given by $e_1' M^{-1} p(|X_i/h|) f_{X,+}^{-1}(Y_i - \mu_+)I(X_i \geq 0)$ and the remaining d_Y columns are given by $e_1' M^{-1} p(|X_i/h|) f_{X,-}^{-1}(Y_i - \mu_-)I(X_i < 0)$. Note that the above expression can be written as

$$\begin{aligned}&T(X_i/h, Y_i, \nu_u(h), \nu_\ell(h), \alpha(h), \{\beta_{u,j,m}(h)\}_{1\le j\le d_Y, 1\le m\le r+1}, \{\beta_{\ell,j,m}(h)\}_{1\le j\le d_Y, 1\le m\le r+1})\\ &- T(X_i/h, Y_i, \nu_u(0), \nu_\ell(0), \alpha(0), \{\beta_{u,j,m}(0)\}_{1\le j\le d_Y, 1\le m\le r+1}, \{\beta_{\ell,j,m}(0)\}_{1\le j\le d_Y, 1\le m\le r+1})\end{aligned}$$

for a function T that is Lipschitz in its remaining arguments uniformly over $X_i/h, Y_i$ on bounded sets. Combining this with the previous lemmas gives the result. □

It follows from Lemmas S2.7 and S1.12 that the conclusion of Lemma S2.7 also holds with $\sigma(h)\psi(W_i, h)$ replaced by $\psi(W_i, h)$, so long as the remaining conditions of Lemma S1.12 (those involving the conditional expectation and variance of $\psi(W_i, 0)$) hold. We have

$$\begin{aligned}&E[\psi(W_i, 0)|X_i = x]\\ &= \frac{1}{\sigma(0)}\left\{D_{g,u}(\alpha(0)) f_{X,+}^{-1}[\tilde{\mu}(x) - \tilde{\mu}_+] I(x \ge 0) + D_{g,\ell}(\alpha(0)) f_{X,-}^{-1}[\tilde{\mu}(x) - \tilde{\mu}_-] I(x < 0)\right\}\end{aligned}$$

and

$$\begin{aligned}var[\psi(W_i, 0)|X_i = x] = \frac{1}{\sigma^2(0)}&\left\{D_{g,u}(\alpha(0))\tilde{\Sigma}(x) D_{g,u}(\alpha(0))' f_{X,+}^{-2} I(x \ge 0)\right.\\ &\left.+ D_{g,\ell}(\alpha(0))\tilde{\Sigma}(x) D_{g,\ell}(\alpha(0))' f_{X,-}^{-2} I(x < 0)\right\}\end{aligned}$$

By the conditions on $\tilde{\mu}(x)$ and $\tilde{\Sigma}(x)$, it follows that these expressions are left and right continuous in x at 0 with modulus $\ell(x)$ satisfying the necessary conditions. By this and the conditions on f_X, it follows that the same holds for $E[\psi(W_i, 0)||X_i| = x]$ and $var[\psi(W_i, 0)||X_i| = x]$. In addition, the assumptions guarantee that $var[\psi(W_i, 0)||X_i| = x]$ is bounded away from zero for small x so that $\sigma(0) > 0$.

Thus, for $\psi(W_i, h)$ defined above,

$$\begin{aligned}
&\sup_{\underline{h}_n \le h \le \overline{h}_n} \left\| \frac{\sqrt{nh}(\hat{\theta}(h) - \theta(h))}{\hat{\sigma}(h)} - \frac{1}{\sqrt{nh}} \sum_{i=1}^{n} \psi(W_i, h) k(X_i/h) \right\| \\
&\le \sup_{\underline{h}_n \le h \le \overline{h}_n} \left\| \frac{\sqrt{nh}(\hat{\theta}(h) - \theta(h))}{\sigma(h)} - \frac{1}{\sqrt{nh}} \sum_{i=1}^{n} \psi(W_i, h) k(X_i/h) \right\| \\
&+ \sup_{\underline{h}_n \le h \le \overline{h}_n} \left\| \sqrt{nh}(\hat{\theta}(h) - \theta(h)) \right\| \cdot \left\| \frac{1}{\sigma(h)} - \frac{1}{\hat{\sigma}(h)} \right\|.
\end{aligned}$$

By Lemma S2.3, the first term is $\mathcal{O}_P\left(1/\sqrt{\log\log \underline{h}_n^{-1}}\right)$, and the last term is $\mathcal{O}_P\left(\sqrt{\log\log \underline{h}_n^{-1}} \cdot \frac{\sqrt{\log\log \underline{h}_n^{-1}}}{\sqrt{n\underline{h}_n}}\right)$. Thus, for $(\log\log \underline{h}_n^{-1})^3/n\underline{h}_n \to 0$, both terms will be $o_P(1/\sqrt{\log\log \underline{h}_n^{-1}})$ as required. This completes the proof of Theorem S2.1.

S2.1 Equivalent Kernels for Local Linear Regression

Thus section computes the equivalent kernels $k(u) = e_1' M^{-1} p(|u|) k^*(u)$ for the local linear estimator ($r = 1$) and the local quadratic estimator ($r = 2$) for some popular choices of the kernel k^*. For $r = 1$, we have

$$e_1' M^{-1} p(u) = e_1' \begin{pmatrix} \mu_{k^*,0} & \mu_{k^*,1} \\ \mu_{k^*,1} & \mu_{k^*,2} \end{pmatrix}^{-1} \begin{pmatrix} 1 \\ |u| \end{pmatrix} = \frac{\mu_{k^*,2} - \mu_{k^*,1}|u|}{\mu_{k^*,0}\mu_{k^*,2} - \mu_{k^*,1}^2}.$$

For $r = 2$, we have

$$e_1 M^{-1} p(u) = \frac{1}{D} \left((\mu_{k^*,4}\mu_{k^*,2} - \mu_{k^*,3}^2) + (\mu_{k^*,1}\mu_{k^*,4} - \mu_{k^*,2}\mu_{k^*,3}) |u| + \left(\mu_{k^*,2}^2 - \mu_{k^*,1}\mu_{k^*,3}\right) u^2 \right),$$

where $D = \det(M) = \mu_{k^*,0}(\mu_{k^*,2}\mu_{k^*,4} - \mu_{k^*,3}^2) - \mu_{k^*,1}(\mu_{k^*,1}\mu_{k^*,4} - \mu_{k^*,2}\mu_{k^*,3}) + \mu_{k^*,2}(\mu_{k^*,1}\mu_{k^*,3} - \mu_{k^*,2}^2)$. The moments $\mu_{k^*,j}$ for the uniform, triangular, and Epanechnikov kernel are given by

Name	μ_0	μ_1	μ_2	μ_3	μ_4
Uniform	$\frac{1}{2}$	$\frac{1}{4}$	$\frac{1}{6}$	$\frac{1}{8}$	$\frac{1}{10}$
Triangular	$\frac{1}{2}$	$\frac{1}{6}$	$\frac{1}{12}$	$\frac{1}{20}$	$\frac{1}{30}$
Epanechnikov	$\frac{1}{2}$	$\frac{3}{16}$	$\frac{1}{10}$	$\frac{1}{16}$	$\frac{3}{70}$

Plugging these moments into the definitions of equivalent kernels in the two displays above then yields the definitions of equivalent kernels for local linear and local quadratic regressions. These definitions are summarized in Table 6.

S3 Applications

This section gives proofs of the results for the applications in the main text.

S3.1 Regression Discontinuity/LATEs for Largest Sets of Compliers

This section proves Theorems 5.1 and 5.2. First, note that the regression discontinuity and LATE applications can both be written as functions of local polynomial estimators in the above setup, with $d_Y = 2$ and Y_i playing the role of $Y_{i,1}$ and D_i playing the role of $Y_{i,2}$. For the LATE application, we define $X_i = -(Z_i - \underline{z})I(|Z_i - \underline{z}| \le |Z_i - \overline{z}|) + (\overline{z} - Z_i)I(|Z_i - \underline{z}| > |Z_i - \overline{z}|)$. Both of these applications fit into the setup of Section S2 with, letting $\alpha(h) = (\alpha_u(h)', \alpha_\ell(h)') = (\alpha_{u,Y}(h), \alpha_{u,D}(h), \alpha_{\ell,Y}(h), \alpha_{\ell,D}(h))'$ (where we use the suggestive subscripts "Y" and "D" rather than 1 and 2), $g(\alpha) = \frac{\alpha_{u,Y} - \alpha_{\ell,Y}}{\alpha_{u,D} - \alpha_{\ell,D}}$. Then, letting $\Delta_D = \alpha_{u,D} - \alpha_{\ell,D}$, we have

$$D_g(\alpha) = \begin{bmatrix} \frac{1}{\Delta_D} & \frac{-g(\alpha)}{\Delta_D} & \frac{-1}{\Delta_D} & \frac{g(\alpha)}{\Delta_D} \end{bmatrix}.$$

This is Lipschitz continuous and bounded over bounded sets with $\alpha_{u,D} - \alpha_{\ell,D}$ bounded away from zero.

For the last condition (nondegeneracy of the conditional variance), note that $D_{g,u}(\alpha(0))\tilde{\Sigma}_+ D_{g,u}(\alpha(0)) = \frac{1}{\Delta_D(0)^2} var[Y_i - g(\alpha(0))D_i | X_i = 0_+]$, which will be nonzero so long as $corr(D_i, Y_i | X_i = 0_+) < 1$ and $var(Y_i | X_i = 0_+) > 0$. A sufficient condition for this is that $var(Y_i | D_i = d, X_i = 0_+) > 0$ is nonzero for $d = 0$ or $d = 1$, and this (or the corresponding statement with $+$ replaced by $-$) holds under the conditions of the theorem.

S3.2 Trimmed Average Treatment Effects under Unconfoundedness

This section proves Theorem 5.3. We first give an intuitive derivation of the critical value, which explains why it differs in this setting, and provide the technical details at the end.

To derive the form of the correction in this case, note that, under the conditions of the theorem, $\frac{\sqrt{n}(\hat{\theta}(h)-\theta(h))}{\hat{\sigma}(h)}$ will converge to a Gaussian process $\mathbb{G}(h)$ with covariance

$$cov(\mathbb{G}(h),\mathbb{G}(h')) = \frac{cov\left\{[\tilde{Y}_i - \theta(h)]I(X_i \in \mathcal{X}_h), [\tilde{Y}_i - \theta(h')]I(X_i \in \mathcal{X}_{h'})\right\}}{\sqrt{var\left\{[\tilde{Y}_i - \theta(h)]I(X_i \in \mathcal{X}_h)\right\} var\left\{[\tilde{Y}_i - \theta(h')]I(X_i \in \mathcal{X}_{h'})\right\}}}.$$

Let $v(h) = var\{[\tilde{Y}_i - \theta(h)]I(X_i \in \mathcal{X}_h)\}$ as defined in the statement of the theorem. Note that, for $h \geq h'$,

$$\begin{aligned} &cov\left\{[\tilde{Y}_i - \theta(h)]I(X_i \in \mathcal{X}_h), [\tilde{Y}_i - \theta(h')]I(X_i \in \mathcal{X}_{h'})\right\} = E\left\{[\tilde{Y}_i - \theta(h)][\tilde{Y}_i - \theta(h')]I(X_i \in \mathcal{X}_h)\right\} \\ &= E\left\{[\tilde{Y}_i - \theta(h)]^2 I(X_i \in \mathcal{X}_h)\right\} + [\theta(h) - \theta(h')]E\left\{[\tilde{Y}_i - \theta(h)]I(X_i \in \mathcal{X}_h)\right\} = v(h) \end{aligned}$$

where the last step follows since $E\left\{[\tilde{Y}_i - \theta(h)]I(X_i \in \mathcal{X}_h)\right\} = 0$. Note also that $v(h)$ is weakly decreasing in h, which can be seen by noting that $v(h) = \inf_a E\left\{[\tilde{Y}_i - a]^2 I(X_i \in \mathcal{X}_h)\right\}$, since $\theta(h)$ is the conditional expectation of $\tilde{Y}_i$ given $X_i \in \mathcal{X}_h$. Thus,

$$cov(\mathbb{G}(h),\mathbb{G}(h')) = \frac{v(h \vee h')}{\sqrt{v(h)v(h')}} = \frac{v(h) \wedge v(h')}{\sqrt{v(h)v(h')}},$$

so $\mathbb{G}(h) \overset{d}{=} \frac{\mathbb{B}(v(h))}{\sqrt{v(h)}}$ where $\mathbb{B}$ is a Brownian motion. Thus, the distribution of $\sup_{\underline{h} \leq h \leq \overline{h}} \frac{\sqrt{n}(\hat{\theta}(h)-\theta(h))}{\hat{\sigma}(h)}$ can be approximated by the distribution of $\sup_{v(\overline{h}) \leq t \leq v(\underline{h})} \frac{\mathbb{B}(t)}{\sqrt{t}} \overset{d}{=} \sup_{1 \leq t \leq v(\underline{h})/v(\overline{h})} \frac{\mathbb{B}(t)}{\sqrt{t}}$. Note that $v(h) = \sigma(h)^2 P(X_i \in \mathcal{X}_h)^2$, so that

$$\frac{v(\underline{h})}{v(\overline{h})} = \frac{\sigma(\underline{h})^2 P(X_i \in \mathcal{X}_{\underline{h}})^2}{\sigma(\overline{h})^2 P(X_i \in \mathcal{X}_{\overline{h}})^2}.$$

Thus, $\hat{t}$ is a consistent estimator for $\frac{v(\underline{h})}{v(\overline{h})}$ under the conditions of the theorem.

To formalize these arguments, note that, by Theorem 19.5 in van der Vaart (1998), $\frac{\sqrt{n}(\hat{\theta}(h)-\theta(h))}{\hat{\sigma}(h)} \overset{d}{\to} \mathbb{G}(h)$, taken as processes over $h \in [\underline{h}, \overline{h}]$ with the supremum norm. By the calculations above,

$$\sup_{h \in [\underline{h}, \overline{h}]} |\mathbb{G}(h)| \overset{d}{=} \sup_{h \in [\underline{h}, \overline{h}]} \left| \frac{\mathbb{B}(v(h))}{\sqrt{v(h)}} \right|$$

where $\mathbb{B}$ is a Brownian motion. The result then follows since $\{t | v(h) = t \text{ some } h \in [\underline{h}, \overline{h}]\} \subseteq$

$[v(\overline{h}), v(\underline{h})]$, and the two sets are equal if $v(h)$ is continuous.

S4 Critical values

Tables S1–S6 give one- and two-sided critical values $c_{1-\alpha}\left(\overline{h}_n/\underline{h}_n, k\right)$ and $c_{1-\alpha,|\cdot|}\left(\overline{h}_n/\underline{h}_n, k\right)$ for several kernel functions k, α and a selected of values of $\overline{h}_n/\underline{h}_n$. The Critical values can also be obtained using our R package `bandwidth-snooping`, which can be downloaded from `https://github.com/kolesarm/bandwidth-snooping`. Tables S1–S2 give one- and two-sided critical values for local constant (Nadaraya-Watson) regression in the interior of the support of the regressor. Tables S3–S4 give one- and two-sided critical values for local linear regression at a boundary. Tables S5–S6 give one- and two-sided critical values for local quadratic regression at a boundary.

Critical values for other choices of $\overline{h}/\underline{h}$ can be obtained using our package `bandwidth-snooping`, which can be downloaded from `https://github.com/kolesarm/bandwidth-snooping`.

S5 Description of variance estimators used in the Monte Carlo study

Given an i.i.d. sample $\{Y_i, X_i\}_{i=1}^n$, the RD estimator is given by the difference between two polynomial linear regressions of order r with the same bandwidth. We consider local linear ($r = 1$) and local quadratic estimators ($r = 2$). To define the estimators, let $p(x) = (1, x, \ldots, x^r)$ denote a polynomial expansion of order r. Let

$$\hat{\beta}_u(h) = \hat{\Gamma}_u(h)^{-1} \sum_{i=1}^n I(X_i \geq 0) k^*(X_i/h) p(|X_i|) Y_i,$$
$$\hat{\beta}_\ell(h) = \hat{\Gamma}_l(h)^{-1} \sum_{i=1}^n I(X_i < 0) k^*(X_i/h) p(|X_i|) Y_i,$$

where k^* is a kernel, and

$$\hat{\Gamma}_u(h) = \sum_i I(X_i \geq 0) k^*(X_i/h) p(|X_i|) p(|X_i|)',$$
$$\hat{\Gamma}_\ell(h) = \sum_i I(X_i < 0) k^*(X_i/h) p(|X_i|) p(|X_i|)'.$$

Then the estimator is given by

$$\hat{\theta}(h) = \hat{\alpha}_u(h) - \hat{\alpha}_\ell(h),$$

where $\hat{\alpha}_u(h) = e_1'\beta_u(h)$ and $\hat{\alpha}_\ell(h) = e_1'\beta_\ell(h)$.

All variance estimators have the form

$$\hat{\sigma}^2(h) = nh\left(\widehat{var}(\hat{\alpha}_u(h)) + \widehat{var}(\hat{\alpha}_\ell(h))\right).$$

Following the recommendation of Imbens and Lemieux (2008), the plug-in estimator sets

$$\widehat{var}(\hat{\alpha}_u(h)) = \frac{\int k^2(u)\,du}{\hat{f}_{X,h}(0)}\hat{\sigma}_u^2,$$

where k is an equivalent kernel, $\hat{f}_{X,h}(0) = \sum_{i=1}^n I(|X|_i \leq h)/(2nh)$, and $\hat{\sigma}_u^2 = \sum_{i=1}^n I(0 \leq |X|_i \leq h)(Y_i - \hat{\alpha}_u(h))^2 / \sum_{i=1}^n I(0 \leq |X|_i \leq h)$. The expression for $\widehat{var}(\hat{\alpha}_\ell(u))$ is similar.

The remaining variance estimators all have the form

$$\widehat{var}(\hat{\alpha}_u(h)) = e_1'\hat{\Gamma}_u(h)^{-1}\left(\sum_{i=1}^n I(X_i \geq 0)\hat{\sigma}_u^2(X_i)k^*(X_i/h)p(|X_i|)p(|X_i|)'\right)\hat{\Gamma}_u(h)^{-1}e_1$$

and similarly for $\widehat{var}(\hat{\alpha}_u(h))$, where $\hat{\sigma}_u^2(X_i)$ and $\hat{\sigma}_\ell^2(X_i)$ are some estimators of $var(Y_i \mid X_i)$. The EHW estimator sets $\hat{\sigma}_u^2(X_i) = (Y_i - X_i'\hat{\beta}_u)^2$. Following the recommendation of Calonico, Cattaneo, and Titiunik (2014), the NN estimator sets

$$\hat{\sigma}_u^2(X_i) = I(X_i \geq 0)\frac{J}{J+1}\left(Y_i - \sum_{j=1}^J Y_{\ell_{u,j}(i)}\right)^2,$$

where $\ell_{u,j}(i)$ is the jth closest unit to i among $\{k \neq i \colon X_k \geq 0\}$, and $J = 3$. Finally, the exact variance estimator sets $\hat{\sigma}_u^2(X_i) = var(Y_i \mid X_i)$.

S6 Additional Simulations

To examine the effects of heteroscedasticity on the performance of our procedure, we considered two additional simulation designs. The DGP in both designs corresponds to Design 1 in the paper, with the exception that

$$\epsilon_i \mid X_i = x \sim \mathcal{N}(0, \sigma^2(x)),$$

where $\sigma^2(x) = 0.1295^2(1 + |x|)^2$ for Design 3, and $\sigma^2(x) = 0.1295^2(1 - |x|)^2$ for Design 4. Tables S9–S10 report empirical coverage of the confidence bands for $\theta(h)$ for these two additional designs. Tables S7–S8 report empirical coverage of the confidence bands for $\theta(0)$. The results are very similar to Design 1 in the paper.

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	Uniform			Triangular			Epanechnikov		
$\overline{h}/\underline{h}$	0.1	0.05	0.01	0.1	0.05	0.01	0.1	0.05	0.01
1.0	1.29	1.65	2.34	1.29	1.65	2.33	1.29	1.65	2.34
1.2	1.59	1.95	2.62	1.34	1.71	2.39	1.36	1.73	2.40
1.4	1.69	2.04	2.70	1.39	1.75	2.43	1.42	1.78	2.45
1.6	1.76	2.10	2.76	1.42	1.78	2.47	1.46	1.81	2.49
1.8	1.81	2.16	2.81	1.45	1.81	2.49	1.49	1.85	2.51
2.0	1.85	2.19	2.84	1.48	1.84	2.51	1.52	1.88	2.54
3.0	1.97	2.31	2.95	1.57	1.91	2.59	1.61	1.97	2.63
4.0	2.04	2.38	3.01	1.62	1.97	2.63	1.67	2.02	2.68
5.0	2.09	2.43	3.04	1.65	2.01	2.67	1.72	2.06	2.71
6.0	2.12	2.46	3.08	1.68	2.03	2.69	1.75	2.09	2.73
7.0	2.15	2.49	3.10	1.70	2.06	2.71	1.77	2.11	2.75
8.0	2.17	2.51	3.12	1.73	2.08	2.72	1.79	2.13	2.77
9.0	2.19	2.52	3.14	1.74	2.09	2.74	1.81	2.15	2.78
10.0	2.21	2.54	3.16	1.75	2.10	2.74	1.82	2.16	2.79
20.0	2.31	2.63	3.23	1.84	2.17	2.81	1.91	2.24	2.87
50.0	2.41	2.71	3.31	1.93	2.26	2.89	2.00	2.33	2.95
100.0	2.47	2.77	3.36	1.98	2.31	2.93	2.06	2.38	2.99

Table S1: Critical values for one-sided tests with levels $\alpha = 0.1$, 0.05, and 0.01, that correspond to $1-\alpha$ quantiles of $\sup_{1\leq h\leq \overline{h}/\underline{h}} \mathbb{H}(h)$.
Nadaraya-Watson estimator with uniform kernel ($k(u) = \frac{1}{2}I(|u| \leq 1)$), triangular kernel ($k(u) = (1-|u|)_+$), or Epanechnikov kernel ($k(u) = 3/4(1-u^2)_+$).

$\overline{h}/\underline{h}$	Uniform 0.1	Uniform 0.05	Uniform 0.01	Triangular 0.1	Triangular 0.05	Triangular 0.01	Epanechnikov 0.1	Epanechnikov 0.05	Epanechnikov 0.01
1.0	1.64	1.96	2.57	1.64	1.96	2.58	1.64	1.96	2.57
1.2	1.93	2.25	2.85	1.70	2.02	2.64	1.71	2.03	2.65
1.4	2.03	2.35	2.94	1.74	2.06	2.68	1.77	2.08	2.70
1.6	2.09	2.42	3.00	1.78	2.09	2.71	1.81	2.12	2.74
1.8	2.14	2.46	3.05	1.81	2.12	2.73	1.84	2.16	2.77
2.0	2.18	2.50	3.08	1.83	2.15	2.75	1.87	2.18	2.80
3.0	2.30	2.61	3.19	1.91	2.23	2.83	1.96	2.27	2.87
4.0	2.38	2.67	3.24	1.96	2.27	2.87	2.01	2.32	2.92
5.0	2.42	2.71	3.28	2.00	2.30	2.90	2.05	2.36	2.95
6.0	2.45	2.74	3.31	2.03	2.33	2.93	2.08	2.39	2.97
7.0	2.48	2.76	3.34	2.05	2.35	2.95	2.11	2.41	2.99
8.0	2.50	2.78	3.36	2.07	2.37	2.96	2.13	2.42	3.01
9.0	2.52	2.80	3.37	2.08	2.38	2.97	2.14	2.44	3.03
10.0	2.53	2.81	3.39	2.09	2.39	2.98	2.16	2.46	3.03
20.0	2.62	2.89	3.45	2.17	2.47	3.04	2.24	2.53	3.10
50.0	2.71	2.98	3.52	2.25	2.54	3.10	2.32	2.60	3.16
100.0	2.77	3.03	3.56	2.30	2.58	3.14	2.37	2.66	3.20

Table S2: Critical values for two-sided tests with levels $\alpha = 0.1$, 0.05, and 0.01, that correspond to $1-\alpha$ quantiles of $\sup_{1\leq h\leq \overline{h}/\underline{h}}|\mathbb{H}(h)|$.
Nadaraya-Watson estimator with uniform kernel ($k(u) = \frac{1}{2}I(|u| \leq 1)$), triangular kernel ($k(u) = (1-|u|)_+$), or Epanechnikov kernel ($k(u) = 3/4(1-u^2)_+$).

	Uniform			Triangular			Epanechnikov		
$\overline{h}/\underline{h}$	0.1	0.05	0.01	0.1	0.05	0.01	0.1	0.05	0.01
1.0	1.28	1.65	2.33	1.28	1.65	2.34	1.28	1.65	2.33
1.2	1.57	1.94	2.61	1.36	1.73	2.41	1.37	1.74	2.42
1.4	1.68	2.04	2.71	1.41	1.78	2.46	1.44	1.81	2.48
1.6	1.75	2.10	2.77	1.46	1.82	2.49	1.49	1.85	2.53
1.8	1.80	2.15	2.82	1.49	1.86	2.52	1.53	1.89	2.56
2.0	1.84	2.19	2.86	1.52	1.89	2.55	1.56	1.92	2.58
3.0	1.97	2.32	2.97	1.62	1.97	2.63	1.67	2.02	2.68
4.0	2.05	2.39	3.02	1.68	2.03	2.68	1.74	2.08	2.74
5.0	2.09	2.43	3.06	1.72	2.07	2.71	1.78	2.12	2.77
6.0	2.13	2.46	3.09	1.75	2.09	2.74	1.81	2.15	2.80
7.0	2.15	2.49	3.12	1.78	2.12	2.76	1.84	2.18	2.81
8.0	2.18	2.51	3.14	1.80	2.14	2.77	1.86	2.20	2.83
9.0	2.20	2.53	3.15	1.81	2.15	2.79	1.88	2.21	2.85
10.0	2.21	2.54	3.16	1.83	2.17	2.81	1.89	2.23	2.86
20.0	2.31	2.63	3.23	1.91	2.24	2.88	1.98	2.31	2.93
50.0	2.41	2.72	3.31	2.00	2.32	2.94	2.07	2.39	3.00
100.0	2.47	2.77	3.36	2.06	2.38	2.98	2.13	2.44	3.04

Table S3: Critical values for one-sided tests with levels $\alpha = 0.1$, 0.05, and 0.01, that correspond to $1-\alpha$ quantiles of $\sup_{1\le h\le \overline{h}/\underline{h}} \mathbb{H}(h)$.
Local linear regression at a boundary with uniform kernel ($k^*(u) = \frac{1}{2}I(|u| \le 1)$), triangular kernel ($k^*(u) = (1-|u|)_+$), or Epanechnikov kernel ($k^*(u) = 3/4(1-u^2)_+$).

	Uniform			Triangular			Epanechnikov		
$\overline{h}/\underline{h}$	0.1	0.05	0.01	0.1	0.05	0.01	0.1	0.05	0.01
1.0	1.64	1.96	2.57	1.64	1.96	2.58	1.64	1.96	2.57
1.2	1.93	2.25	2.86	1.72	2.04	2.65	1.74	2.05	2.67
1.4	2.03	2.34	2.96	1.77	2.09	2.70	1.80	2.11	2.73
1.6	2.10	2.40	3.01	1.81	2.12	2.74	1.84	2.16	2.77
1.8	2.15	2.46	3.06	1.85	2.16	2.77	1.88	2.19	2.81
2.0	2.19	2.49	3.09	1.87	2.19	2.80	1.91	2.22	2.84
3.0	2.31	2.61	3.19	1.97	2.28	2.87	2.01	2.32	2.91
4.0	2.38	2.68	3.24	2.02	2.33	2.92	2.07	2.38	2.96
5.0	2.42	2.72	3.29	2.06	2.36	2.95	2.11	2.41	2.99
6.0	2.45	2.75	3.31	2.09	2.39	2.97	2.14	2.44	3.02
7.0	2.48	2.77	3.33	2.11	2.41	2.99	2.17	2.46	3.03
8.0	2.50	2.79	3.35	2.13	2.43	3.00	2.19	2.48	3.05
9.0	2.52	2.81	3.36	2.15	2.44	3.01	2.20	2.50	3.06
10.0	2.54	2.82	3.37	2.16	2.45	3.03	2.22	2.51	3.08
20.0	2.62	2.90	3.44	2.24	2.53	3.09	2.30	2.58	3.13
50.0	2.72	2.98	3.52	2.32	2.60	3.16	2.39	2.66	3.21
100.0	2.77	3.04	3.56	2.37	2.65	3.20	2.44	2.71	3.25

Table S4: Critical values for two-sided tests with levels $\alpha = 0.1$, 0.05, and 0.01, that correspond to $1-\alpha$ quantiles of $\sup_{1\leq h\leq \overline{h}/\underline{h}}|\mathbb{H}(h)|$.
Local linear regression at a boundary with uniform kernel ($k^*(u) = \frac{1}{2}I(|u| \leq 1)$), triangular kernel ($k^*(u) = (1-|u|)_+$), or Epanechnikov kernel ($k^*(u) = 3/4(1-u^2)_+$).

	Uniform			Triangular			Epanechnikov		
$\overline{h}/\underline{h}$	0.1	0.05	0.01	0.1	0.05	0.01	0.1	0.05	0.01
1.0	1.28	1.65	2.34	1.29	1.66	2.33	1.29	1.66	2.32
1.2	1.58	1.95	2.61	1.38	1.75	2.42	1.40	1.76	2.44
1.4	1.69	2.05	2.71	1.44	1.80	2.47	1.47	1.83	2.49
1.6	1.76	2.11	2.76	1.49	1.85	2.52	1.52	1.88	2.54
1.8	1.81	2.17	2.82	1.53	1.89	2.55	1.56	1.92	2.58
2.0	1.85	2.21	2.85	1.56	1.92	2.58	1.60	1.95	2.61
3.0	1.98	2.32	2.95	1.66	2.01	2.67	1.71	2.06	2.71
4.0	2.05	2.39	3.01	1.72	2.07	2.71	1.78	2.12	2.75
5.0	2.10	2.43	3.04	1.76	2.10	2.74	1.82	2.16	2.79
6.0	2.13	2.47	3.07	1.79	2.13	2.77	1.85	2.19	2.82
7.0	2.16	2.49	3.09	1.82	2.16	2.80	1.88	2.21	2.84
8.0	2.18	2.51	3.12	1.84	2.17	2.81	1.90	2.23	2.86
9.0	2.20	2.53	3.14	1.86	2.19	2.83	1.92	2.25	2.88
10.0	2.22	2.54	3.15	1.87	2.21	2.84	1.93	2.27	2.89
20.0	2.31	2.63	3.23	1.96	2.28	2.92	2.02	2.34	2.97
50.0	2.40	2.71	3.31	2.04	2.37	2.99	2.11	2.43	3.04
100.0	2.46	2.77	3.36	2.10	2.42	3.03	2.16	2.48	3.09

Table S5: Critical values for one-sided tests with levels $\alpha = 0.1$, 0.05, and 0.01, that correspond to $1-\alpha$ quantiles of $\sup_{1 \leq h \leq \overline{h}/\underline{h}} \mathbb{H}(h)$.
Local quadratic regression at a boundary with uniform kernel ($k^*(u) = \frac{1}{2} I(|u| \leq 1)$), triangular kernel ($k^*(u) = (1-|u|)_+$), or Epanechnikov kernel ($k^*(u) = 3/4(1-u^2)_+$).

	Uniform			Triangular			Epanechnikov		
$\overline{h}/\underline{h}$	0.1	0.05	0.01	0.1	0.05	0.01	0.1	0.05	0.01
1.0	1.65	1.96	2.58	1.65	1.96	2.58	1.65	1.96	2.58
1.2	1.94	2.25	2.86	1.74	2.05	2.67	1.75	2.06	2.68
1.4	2.04	2.34	2.94	1.79	2.11	2.72	1.82	2.13	2.74
1.6	2.10	2.41	3.00	1.84	2.15	2.76	1.87	2.18	2.78
1.8	2.15	2.46	3.06	1.88	2.18	2.80	1.92	2.23	2.82
2.0	2.19	2.50	3.09	1.91	2.22	2.83	1.95	2.25	2.86
3.0	2.31	2.61	3.21	2.00	2.31	2.91	2.05	2.35	2.96
4.0	2.38	2.67	3.26	2.06	2.36	2.95	2.11	2.41	3.00
5.0	2.42	2.71	3.30	2.10	2.40	2.99	2.15	2.44	3.03
6.0	2.46	2.74	3.32	2.12	2.42	3.01	2.18	2.47	3.06
7.0	2.48	2.77	3.35	2.15	2.44	3.03	2.20	2.50	3.08
8.0	2.50	2.79	3.37	2.16	2.46	3.04	2.22	2.51	3.10
9.0	2.52	2.80	3.38	2.18	2.48	3.05	2.24	2.53	3.11
10.0	2.54	2.82	3.40	2.20	2.49	3.07	2.25	2.54	3.12
20.0	2.63	2.90	3.45	2.28	2.56	3.14	2.34	2.62	3.19
50.0	2.71	2.98	3.53	2.36	2.64	3.21	2.42	2.70	3.25
100.0	2.76	3.03	3.56	2.41	2.69	3.24	2.47	2.75	3.29

Table S6: Critical values for two-sided tests with levels $\alpha = 0.1$, 0.05, and 0.01, that correspond to $1-\alpha$ quantiles of $\sup_{1\leq h\leq \overline{h}/\underline{h}}|\mathbb{H}(h)|$.
Local quadratic regression at a boundary with uniform kernel ($k(u) = \frac{1}{2}I(|u| \leq 1)$), triangular kernel ($k(u) = (1-|u|)_+$), or Epanechnikov kernel ($k(u) = 3/4(1-u^2)_+$).

$(\underline{h}, \bar{h})$	$\hat{\sigma}(h)$	Uniform Kernel Pointwise	Uniform Kernel Naive	Uniform Kernel Adjusted	Triangular Kernel Pointwise	Triangular Kernel Naive	Triangular Kernel Adjusted
Local Linear regression							
$(1/2\hat{h}_{IK}, \hat{h}_{IK})$	exact	(78.9, 90.7)	63.9	82.6	(80.4, 91.0)	76.8	83.5
	EHW	(77.9, 89.4)	62.2	81.1	(79.2, 89.2)	75.0	81.8
	plugin	(84.7, 92.0)	70.0	86.3	(90.3, 93.8)	86.2	90.8
	NN	(81.0, 91.3)	67.0	84.2	(82.1, 91.5)	78.8	85.0
$(1/2\hat{h}_{IK}, 2\hat{h}_{IK})$	exact	(78.8, 91.1)	54.6	82.7	(80.4, 91.0)	71.9	83.7
	EHW	(77.9, 90.5)	53.0	81.0	(79.2, 89.2)	70.2	81.9
	plugin	(84.7, 97.6)	65.0	87.7	(90.3, 98.3)	85.2	92.5
	NN	(80.9, 92.7)	58.7	84.8	(82.1, 91.5)	74.8	85.5
$(1/4\hat{h}_{IK}, 1/2\hat{h}_{IK})$	exact	(91.2, 95.0)	82.4	94.0	(91.4, 95.0)	88.7	92.9
	EHW	(89.8, 92.6)	78.4	91.4	(89.7, 92.0)	84.8	89.6
	plugin	(92.4, 95.9)	85.7	95.2	(94.1, 96.5)	92.0	95.0
	NN	(91.8, 94.6)	83.1	93.9	(91.9, 94.1)	88.4	92.3
Local quadratic regression							
$(1/2\hat{h}_{IK}, \hat{h}_{IK})$	NN	(92.1, 95.1)	82.5	93.6	(91.1, 94.7)	87.6	92.2
$(1/2\hat{h}_{IK}, 2\hat{h}_{IK})$	NN	(83.7, 95.1)	60.2	86.3	(85.0, 94.7)	76.1	87.1
$(1/4\hat{h}_{IK}, 1/2\hat{h}_{IK})$	NN	(93.9, 94.8)	84.4	94.1	(93.1, 94.3)	88.7	92.9

Table S7: Monte Carlo study of regression discontinuity. Design 3. Empirical coverage of $\theta(0)$ for nominal 95% confidence bands around IK bandwidth. "Pointwise" refers to range of coverages of pointwise confidence intervals. "Naive" refers to the coverage of the naive confidence band that uses the unadjusted critical value equal to 1.96. "Adjusted" refers to confidence bands using adjusted critical values based on Theorem 3.1. Variance estimators are described in the text. 50,000 Monte Carlo draws (10,000 for NN-based variance estimators), 100 grid points for h.

$(\underline{h}, \bar{h})$	$\hat{\sigma}(h)$	Uniform Kernel Pointwise	Uniform Kernel Naive	Uniform Kernel Adjusted	Triangular Kernel Pointwise	Triangular Kernel Naive	Triangular Kernel Adjusted
Local Linear regression							
$(1/2\hat{h}_{IK}, \hat{h}_{IK})$	exact	(70.8, 91.8)	63.9	82.8	(76.1, 92.2)	75.0	82.1
	EHW	(69.5, 90.3)	62.2	80.8	(74.4, 90.2)	72.9	79.9
	plugin	(72.5, 90.5)	63.6	80.5	(78.3, 90.4)	75.9	81.9
	NN	(73.5, 92.6)	67.0	84.5	(78.1, 92.3)	77.1	83.5
$(1/2\hat{h}_{IK}, 2\hat{h}_{IK})$	exact	(69.3, 91.8)	57.5	83.2	(75.1, 92.2)	71.1	82.8
	EHW	(68.1, 90.3)	56.0	81.2	(73.5, 90.2)	69.2	80.8
	plugin	(72.1, 98.1)	60.2	82.3	(78.3, 97.6)	75.3	84.5
	NN	(72.1, 92.6)	61.1	85.0	(77.2, 92.3)	73.8	84.4
$(1/4\hat{h}_{IK}, 1/2\hat{h}_{IK})$	exact	(92.3, 95.2)	85.6	95.5	(92.7, 95.2)	90.2	94.1
	EHW	(90.8, 92.9)	81.5	92.7	(90.6, 92.3)	86.2	90.5
	plugin	(91.1, 95.4)	85.2	94.6	(90.9, 95.3)	89.0	92.8
	NN	(93.0, 94.8)	85.9	94.7	(92.8, 94.2)	89.7	93.0
Local quadratic regression							
$(1/2\hat{h}_{IK}, \hat{h}_{IK})$	NN	(93.0, 95.3)	86.9	95.5	(92.3, 94.9)	89.4	93.6
$(1/2\hat{h}_{IK}, 2\hat{h}_{IK})$	NN	(72.6, 95.3)	63.8	86.6	(78.9, 94.9)	74.4	85.6
$(1/4\hat{h}_{IK}, 1/2\hat{h}_{IK})$	NN	(93.6, 94.7)	85.8	94.5	(93.1, 94.3)	89.1	93.2

Table S8: Monte Carlo study of regression discontinuity. Design 4. Empirical coverage of $\theta(0)$ for nominal 95% confidence bands around IK bandwidth. "Pointwise" refers to range of coverages of pointwise confidence intervals. "Naive" refers to the coverage of the naive confidence band that uses the unadjusted critical value equal to 1.96. "Adjusted" refers to confidence bands using adjusted critical values based on Theorem 3.1. Variance estimators are described in the text. 50,000 Monte Carlo draws (10,000 for NN-based variance estimators), 100 grid points for h.

$(\underline{h}, \bar{h})$	$\hat{\sigma}(h)$	Uniform Kernel Pointwise	Uniform Kernel Naive	Uniform Kernel Adjusted	Triangular Kernel Pointwise	Triangular Kernel Naive	Triangular Kernel Adjusted
Local Linear regression							
$(1/2\hat{h}_{IK}, \hat{h}_{IK})$	exact	(94.7, 95.1)	84.4	95.0	(94.7, 95.1)	91.2	94.7
	EHW	(94.0, 94.4)	82.6	93.9	(93.6, 94.3)	89.5	93.4
	plugin	(96.5, 97.6)	90.2	97.2	(97.4, 98.7)	96.1	97.9
	NN	(95.4, 95.9)	85.9	95.7	(95.2, 95.8)	91.9	94.9
$(1/2\hat{h}_{IK}, 2\hat{h}_{IK})$	exact	(91.1, 95.1)	72.0	93.2	(92.9, 95.1)	86.0	93.6
	1residu	(95.1, 97.9)	81.5	96.5	(94.2, 96.5)	87.7	94.3
	EHW	(90.6, 94.4)	69.8	92.0	(92.3, 94.3)	84.3	92.3
	plugin	(96.5, 99.2)	87.7	98.0	(97.4, 99.7)	95.8	98.4
	NN	(92.7, 95.9)	74.5	93.8	(93.9, 95.8)	87.6	94.2
$(1/4\hat{h}_{IK}, 1/2\hat{h}_{IK})$	exact	(95.0, 95.3)	86.1	95.9	(95.1, 95.3)	91.7	95.1
	EHW	(92.6, 94.0)	82.2	93.3	(91.8, 93.6)	88.0	92.0
	plugin	(96.4, 96.6)	90.6	97.4	(97.0, 97.3)	95.0	97.1
	NN	(94.7, 95.8)	86.3	95.0	(94.4, 95.2)	91.0	94.0
Local quadratic regression							
$(1/2\hat{h}_{IK}, \hat{h}_{IK})$	NN	(94.9, 95.7)	85.0	95.0	(94.4, 95.3)	90.7	94.3
$(1/2\hat{h}_{IK}, 2\hat{h}_{IK})$	NN	(88.7, 96.3)	71.7	91.9	(92.9, 95.9)	83.5	92.2
$(1/4\hat{h}_{IK}, 1/2\hat{h}_{IK})$	NN	(93.9, 95.0)	84.5	94.1	(93.1, 94.4)	88.7	93.0

Table S9: Monte Carlo study of regression discontinuity. Design 3. Empirical coverage of $\theta(h)$ for nominal 95% confidence bands around IK bandwidth. "Pointwise" refers to range of coverages of pointwise confidence intervals. "Naive" refers to the coverage of the naive confidence band that uses the unadjusted critical value equal to 1.96. "Adjusted" refers to confidence bands using adjusted critical values based on Theorem 3.1. Variance estimators are described in the text. 50,000 Monte Carlo draws (10,000 for NN-based variance estimators), 100 grid points for h.

$(\underline{h}, \bar{h})$	$\hat{\sigma}(h)$	Uniform Kernel			Triangular Kernel		
		Pointwise	Naive	Adjusted	Pointwise	Naive	Adjusted
Local Linear regression							
$(1/2\hat{h}_{IK}, \hat{h}_{IK})$	exact	(94.7, 96.1)	89.3	96.8	(94.7, 96.0)	92.9	95.8
	EHW	(94.1, 95.1)	87.6	95.9	(93.8, 94.7)	91.3	94.5
	plugin	(95.8, 96.4)	90.8	97.1	(95.9, 96.2)	93.8	96.2
	NN	(95.1, 96.2)	89.8	97.0	(95.0, 95.8)	93.0	95.7
$(1/2\hat{h}_{IK}, 2\hat{h}_{IK})$	exact	(91.2, 96.1)	83.2	96.2	(92.7, 96.0)	89.5	95.3
	EHW	(90.7, 95.1)	81.6	95.2	(91.9, 94.7)	87.9	93.9
	plugin	(95.6, 99.0)	89.0	97.4	(95.9, 99.6)	93.4	96.9
	NN	(91.3, 96.2)	83.1	95.5	(93.1, 95.8)	89.3	94.8
$(1/4\hat{h}_{IK}, 1/2\hat{h}_{IK})$	exact	(95.3, 95.9)	88.8	96.8	(95.4, 95.8)	92.8	95.8
	EHW	(92.6, 94.8)	84.7	94.3	(91.9, 94.4)	89.0	92.6
	plugin	(96.1, 96.3)	90.9	97.5	(96.1, 96.2)	93.7	96.3
	NN	(94.6, 96.0)	88.1	95.7	(94.1, 95.6)	91.4	94.5
Local quadratic regression							
$(1/2\hat{h}_{IK}, \hat{h}_{IK})$	NN	(94.7, 96.0)	88.7	96.1	(94.4, 95.7)	92.0	95.1
$(1/2\hat{h}_{IK}, 2\hat{h}_{IK})$	NN	(89.5, 96.5)	81.0	94.9	(90.1, 96.1)	85.3	92.5
$(1/4\hat{h}_{IK}, 1/2\hat{h}_{IK})$	NN	(93.7, 94.7)	85.7	94.6	(93.2, 94.3)	89.3	93.2

Table S10: Monte Carlo study of regression discontinuity. Design 4. Empirical coverage of $\theta(h)$ for nominal 95% confidence bands around IK bandwidth. "Pointwise" refers to range of coverages of pointwise confidence intervals. "Naive" refers to the coverage of the naive confidence band that uses the unadjusted critical value equal to 1.96. "Adjusted" refers to confidence bands using adjusted critical values based on Theorem 3.1. Variance estimators are described in the text. 50,000 Monte Carlo draws (10,000 for NN-based variance estimators), 100 grid points for h.

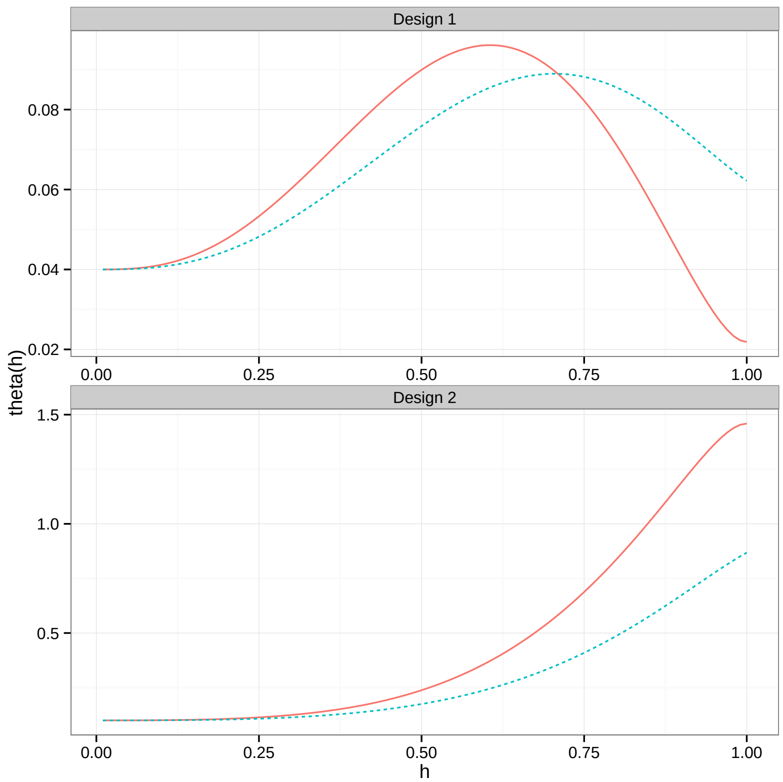

Figure S1: Regression Discontinuity. The function $\theta(h)$ for designs we consider corresponding to local quadratic regression.

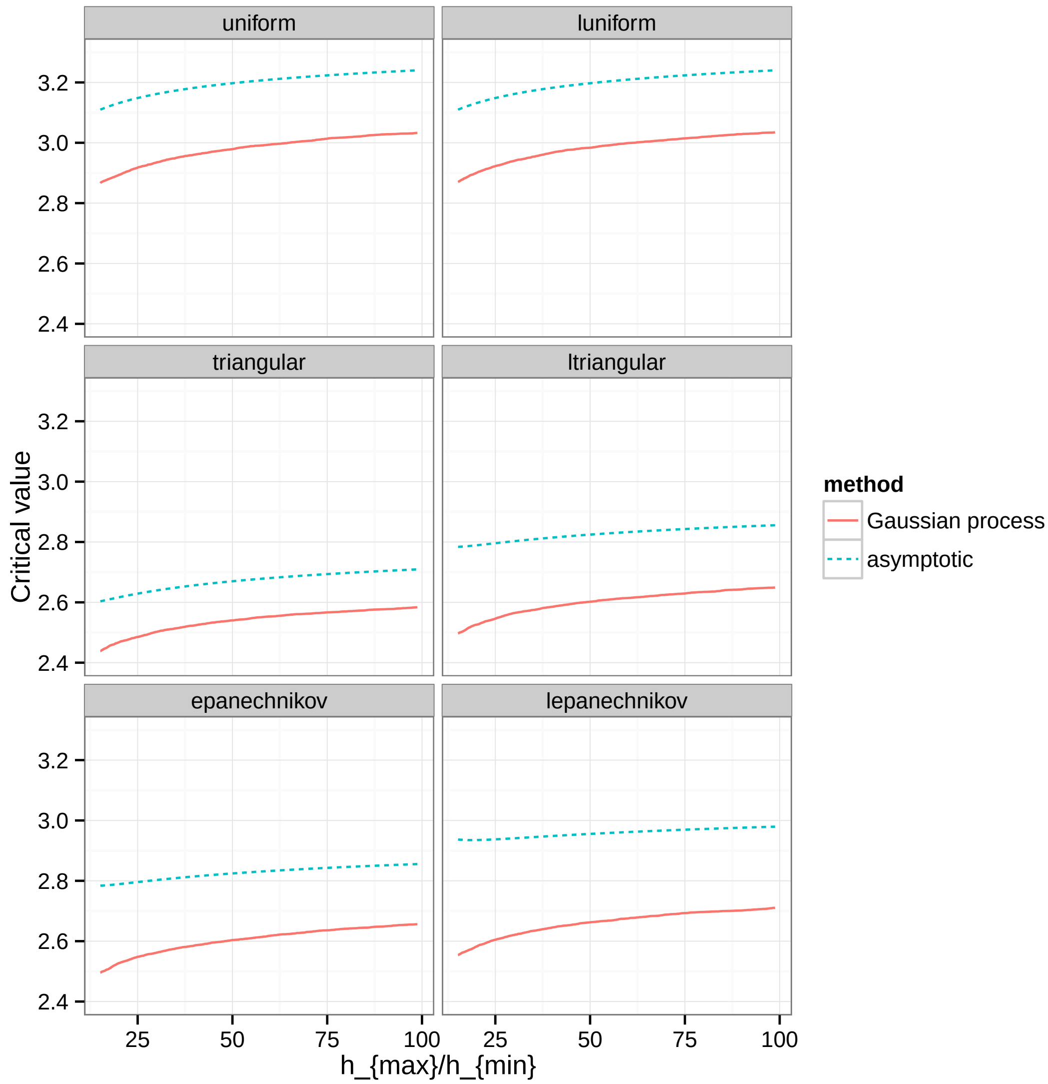

Figure S2: Comparison of critical values based on Gassuan approximation and extreme value approximation (i.e. asymptotic approximation as $\overline{h}/\underline{h} \to \infty$).