One for all and all for one: Dimension reduction for regression checks

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Abstract

We develop a novel dimension-reduction approach to consistent checks of parametric regression models when many regressors are present. The principle is to replace the nonparametric alternative by a class of semiparametric alternatives, namely single-index models, that is rich enough to allow detection of any nonparametric alternative. We propose an omnibus test based on the kernel method that performs against a sequence of directional local nonparametric alternatives as if there was one regressor only, whatever the number of regressors. This test can also be viewed as an improved integrated conditional moment test. A simulation study reveals that our test performs better than several known lack-of-fit tests in multidimensional settings. Our test is little sensitive to the smoothing parameter choice. Moreover, qualitative information can be easily incorporated in the procedure to further improve its power.

Keywords: Dimensionality, Hypothesis testing, Nonparametric methods.

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1 Introduction

Parametric forms are frequently used in regression models to estimate the association between a response variable and predictors. Checking the adequacy of a parametric regression function is then useful in many applications, whether in econometrics or in other applied fields. Popular graphical displays of residuals against covariates can fail to detect an inadequate model when many covariates are present. Hence, since the end of the eighties, many tests for regression checking have been developed. With few exceptions, notably Bierens (1982, 1990) and Stute, Gonzalez Manteiga and Presedo Quindimil (1998), most regression checks rely on some smoothing method, such as kernels, splines, local polynomials, or orthogonal series, from the earlier work of Cox and al. (1988), Azzalini, Bowman and Härdle (1989), Eubank and Spiegleman (1990), Hart and Wehrly (1992), Eubank and Hart (1993), to the more recent papers by Dette (1999), Aerts, Claeskens and Hart (1999), Spokoiny (2001), Baraud, Huet and Laurent (2003). The nice monograph by Hart (1997) reviews this statistical literature, but almost exclusively deals with the one predictor case. Among the authors who explicitly studied the many regressors case, Härdle and Mammen (1993) used an L^2 distance between the parametric regression and the nonparametric one; Zheng (1996), Aerts, Claeskens and Hart (1999), and Guerre and Lavergne (2005) used a score approach; Fan, Zhang and Zhang (2001) adopted a likelihood-ratio approach. The ability of these omnibus tests to detect deviations from the parametric model quickly vanes when there is more than a handful of regressors. Indeed, since the nonparametric estimators suffer from the "curse of dimensionality" as shown by Stone (1980), so too do the related tests. Hence, their usefulness is questionable for many applications, in particular in econometrics where the number of covariates can be large. To circumvent this issue, one can aim at testing the parametric regression against some non-saturated semi-parametric alternatives. Fan, Zhang and Zhang (2001) studied varying coefficients linear models. Aerts, Claeskens and Hart (2000) and Guerre and Lavergne (2005) proposed tests tailored for additive alternatives. Hart (1997, Section 9.3) considered alternatives of the form $m(t(X))$, where $m(\cdot)$ is nonparametric and $t(X)$ is the vector of the first principal components of the covariance matrix X; he noted that there is however no guarantee that lack-of-fit will manifest itself along principal components. Fan and Huang (2001) similarly relied on scores from principal components analysis. The alternative dimension-reduction test of Zhu (2003) assumes *independence* of the parametric residuals with the regressors. All these proposals thus rely on some auxiliary assumptions that allow to restrict the alternative model but yield tests that are not omnibus.

Our goal is to device a powerful regression check that researchers could confidently apply in the presence of many regressors without imposing restrictions on the form of the alternative. In this aim, we develop a novel approach of dimension reduction. Using this approach, we construct a test that improves on known regression checks based on smoothing methods, but can also be viewed as a further elaboration of the integrated conditional moment (ICM) test proposed by Bierens (1982). Our theoretical results show that our test is consistent against any alternative, yet it is not affected by the dimension of the regressors, since it behaves as if there was only one regressor. In practice, we found that our test is more powerful than known lack-of-fit tests in multidimensional settings. Specifically, it outperforms not only the kernel-based test of Zheng (1996), but also the Integrated conditional moment (ICM) test by Bierens (1982) and the test recently proposed Escanciano (2006). Moreover, our test is very little sensitive to the smoothing parameter choice. Finally, we show that one can easily incorporate some qualitative information in the procedure to further improve its power.

Acknowledging that testing directly against saturated alternatives yield low power, our key principle is to replace the nonparametric alternative by a *class* of a semiparametric alternatives that is rich enough to allow detection of any nonparametric alternative, thus reducing the dimension of the problem while preserving consistency. Specifically, we look at the class of single-index regression models. Formally, let $(Y_1, X_1')'$, $\ldots (Y_n, X_n')'$ be independent observations from a population $(Y, X')' \in \mathbb{R}^{1+q}$, where X is a continuous random vector. We want to check whether the regression function $\mathbb{E}(Y|X)$ belongs to a parametric family $\{\mu(\cdot, \theta) : \theta \in \Theta\}$, for instance of linear or logistic functions. Our null hypothesis then writes

$$H_0 \ : \mathbb{E}\left[Y - \mu(X, \theta_0)|X\right] = 0 \quad \text{for some } \theta_0 \ . \tag{1.1}$$

As we face the "curse of dimensionality" in estimating the above conditional expectation, the resulting estimate will be imprecise for even a moderate q, and the related test will lack power. Our approach is instead to estimate conditional expectations given a linear index $X'\beta$ for any β and thus to replace *one* conditional expectation given *all* the regressors by *all* conditional expectations given *one* single linear index only. The advantage is that each expectation can be estimated accurately for a reasonable sample size since it depends on a single linear index only. The apparent drawback is that we have to estimate many conditional expectations. However, this cumbersome task can be avoided by combining expectations into a single integral and estimating this integral at once. We show indeed below that testing H_0 is equivalent to test that

$$\int_{\mathbb{S}^q} \mathbb{E}\left[\mathbb{E}^2\left(Y-\mu(X,\theta_0)|X'\beta\right) f_\beta(X'\beta)\right] d\beta = 0 \quad \text{for some } \theta_0 \,, \tag{1.2}$$

where $\mathbb{S}^q$ is the hypersphere $\{\beta \in \mathbb{R}^q : \|\beta\| = 1\}$ and $f_\beta(\cdot)$ is the density of the linear index $X'\beta$. Our approach thus reduces the dimension of the problem without any knowledge about the form of the alternatives. The resulting test is truly omnibus and the rate of convergence of the test statistic under H_0 equals the rate one would obtain in the one-dimensional case. Moreover, it behaves against local directional alternatives as if there was one regressor only. We also show that when the regressors are bounded, it is sufficient to consider the above integral on a subset of the hypersphere with nonempty interior in (1.2). This allows to incorporate some qualitative information in the procedure. For instance, if it is known that the marginal effects of two regressors X_1 and X_2 always have the same sign, one can choose B as the domain where the corresponding β_1 and β_2 have the same sign.

The rest of the paper is organized as follows. Section 2.1 proves the fundamental lemma on which our dimension-reduction approach relies. Section 2.2 proposes a test statistic based on the kernel method. Section 2.3 deals with the asymptotic behavior of the test under the null hypothesis and its consistency under a sequence of directional alternatives. Section 2.4 justifies the validity of a bootstrap method that can be used instead of the asymptotic approximation to obtain critical values for samples of small or moderate size. Section 3.1 details the practical computation of the test statistic. Section 3.2 reports the results of an extensive simulation study that compare our approach to different tests previously proposed in the literature. The technical proofs are gathered in the Appendix.

2 Methods and results

2.1 The dimension-reduction principle

The following lemma is the crux of our dimension-reduction approach. It provides a direct justification for considering *all* conditional expectations given *one* single linear index for testing H_0. Part (ii) shows that when X is bounded, it is even sufficient to consider *infinitely many* of these conditional expectations. Note that X is bounded can be assumed without loss of generality, since we can always find a one-to-one transformation that maps X in a bounded set and retains all conditioning information.

Lemma 2.1 *Let $\mathbb{S}^q = \{\beta \in \mathbb{R}^q : \|\beta\| = 1\}$ be the hypersphere with radius one. Consider random vectors $Z \in \mathbb{R}$ with $\mathbb{E}(Z^2) < \infty$ and $X \in \mathbb{R}^q$ with bounded density $f(\cdot)$. Let $f_\beta(\cdot)$ be the density of $X'\beta$ and assume that for some C, $|f_\beta(\cdot)| \leq C$ for any $\beta \in \mathbb{S}^q$.*
(i) $\mathbb{E}(Z \mid X) = 0$ is equivalent to

$$\int_{\mathbb{S}^q} \mathbb{E}\left[\mathbb{E}^2\left(Z|X'\beta\right) f_\beta(X'\beta)\right] \, d\beta = 0 \; . \tag{2.3}$$

(ii) If X is bounded, then $\mathbb{E}(Z \mid X) = 0$ is equivalent to

$$\int_{B} \mathbb{E}\left[\mathbb{E}^2\left(Z|X'\beta\right) f_\beta(X'\beta)\right] \, d\beta = 0 \tag{2.4}$$

for any $B \subset \mathbb{S}^q$ with nonempty interior.

As a consequence, our null hypothesis can be (1.1) can be written as (1.2). Moreover, when X is bounded, we can incorporate some qualitative information by considering only a restricted subset B, as mentioned above.

The ICM test of Bierens (1982) and Bierens and Ploberger (1997) is based on the fact that for X bounded, $\mathbb{E}(Z|X) = 0$ iff

$$\int_{\mathbb{R}^q} \left|\mathbb{E}\left[Z\psi\left(X'u\right)\right]\right|^2 \, d\mu(u) = 0 \; ,$$

for well-chosen function $\psi(\cdot)$, such as $\exp(-i\cdot)$ and probability measure $\mu(\cdot)$. As is clear from Lemma 2.1, our approach is related, but instead of choosing a particular $\psi(\cdot)$ at the outset, we pick for each β the function of $X'\beta$ maximizing squared correlation with Z. In so doing, we expect to improve the power of the testing procedure. It is easily shown that the solution is $\mathbb{E}(Z \mid X'\beta)$, which yields our formulation. A similar reasoning applies if one maximizes $\mathbb{E}\left[Z\psi\left(X'\beta\right) f_\beta(X'\beta)\right]$.

Lemma 2.1 can be viewed as deriving from the results of Bierens (1982), but as our lemma is central to our work, we provide here a simple proof and we comment it thereafter.
Proof. (i) The implication is straightforward. By elementary properties of the conditional expectation, for any $\beta \in \mathbb{S}^q$ and any $t \in \mathbb{R}$

$$\psi_\beta(t) := \mathbb{E}\left[\exp\{itX'\beta\}\mathbb{E}(Z \mid X'\beta)\right] = \mathbb{E}\left[\exp\{itX'\beta\}\mathbb{E}(Z \mid X)\right] \; , \tag{2.5}$$

where $i = \sqrt{-1}$. For any $\beta \in \mathbb{S}^q$, $\mathbb{E}\left(Z|X'\beta\right) f_\beta(X'\beta) \in L^1(\mathbb{R}) \cap L^2(\mathbb{R})$, and Parseval's formula yields, see e.g. Rudin (1987),

$$\int_{\mathbb{R}} |\psi_\beta(t)|^2 \, dt = 2\pi \; \mathbb{E}\left[\mathbb{E}^2\left(Z|X'\beta\right) f_\beta(X'\beta)\right]$$

$$\text{and} \quad \int_{\mathbb{S}^q} \int_{\mathbb{R}} |\psi_\beta(t)|^2 \, dt \, d\beta = 2\pi \int_{\mathbb{S}^q} \mathbb{E}^2\left[\mathbb{E}\left(Z|X'\beta\right) f_\beta(X'\beta)\right] \, d\beta \, .$$

If this integral equals zero, this implies $\psi_\beta(t) = 0$ for all β and all t. By the unicity of the Fourier transform, $\mathbb{E}(Z \mid X) = 0$.

(ii) Clearly, $\mathbb{E}(Z \mid X) = 0$ implies (2.4). The equivalence for $B = \mathbb{S}^q$ follows from Part (i). Since

$$2\pi \int_B \mathbb{E}\left[\mathbb{E}^2\left(Z|X'\beta\right) f_\beta(X'\beta)\right] \, d\beta = \int_B \int_{\mathbb{R}} |\psi_\beta(t)|^2 \, dt \, d\beta \, ,$$

(2.4) implies $\psi_\beta(t) = 0$ for all $\beta \in B$ and t. Since X is bounded, this yields $\mathbb{E}(Z|X) = 0$ by Theorem 1 of Bierens (1982). ■

The proof clearly shows how (2.3) naturally appears from Fourier analysis. It is also useful to see that, because of the symmetry of the Fourier transform, our lemma holds not only for the hypersphere $\mathbb{S}^q$, but for any half-hypersphere. By half-hypersphere, we mean any subset H of $\mathbb{S}^q$ such that (i) $H \cup H^- = \mathbb{S}^q$, where $H^- = \{\beta^- : \beta^- = -\beta, \beta \in H\}$ and (ii) $H \cap H^-$ has Lebesgue measure zero. Hence, the assumption of a bounded X is necessary for Part (ii) only if B is not included in a half-hypersphere.

2.2 The test

Let $(Y_i, X_i')'$, $i = 1, \ldots n$, be a random sample from $(Y, X')' \in \mathbb{R}^{1+q}$. The vector X is assumed to be continuously distributed, since regressors with fixed discrete support have no theoretical influence on the asymptotic power of a regression check. The model to be checked writes

$$Y = \mu(X, \theta_0) + \varepsilon, \quad \mathbb{E}\left(\varepsilon|X\right) = 0 \, .$$

An estimated candidate $\widehat{\theta}_n$ for the parameter θ_0 can be obtained by least-squares. The parametric residuals are then $\widehat{U}_i = Y_i - \mu(X_i, \widehat{\theta}_n), i = 1, \ldots n$. We use the kernel method to estimate (2.4), as it yields a very tractable statistic. We could certainly accommodate for other nonparametric methods, such as splines, local polynomials, or orthogonal series, but we do not pursue this issue here. We first define

$$\frac{1}{n(n-1)} \sum_{j \neq i} \widehat{U}_i \widehat{U}_j \frac{1}{h} K_h\left((X_i - X_j)'\beta\right) \, , \tag{2.6}$$

as an estimator of $\mathbb{E}\left[\mathbb{E}^2\left(Y-\mu(X,\theta_0)|X'\beta\right)f_\beta(X'\beta)\right]$. Here $K_h\left(\cdot\right)=K\left(\cdot/h\right)$, where $K(\cdot)$ is an univariate symmetric density and h a bandwidth. This statistic is the one studied by Zheng (1996) and Li and Wang (1998) applied to the index $X'\beta$ and has an asymptotic centered normal distribution with rate $nh^{1/2}$ under H_0. As noted by Dette (1999), Zheng's statistic is comparable to Härdle and Mammen's one (1993) with weight function equal to the squared density, which is exactly what is needed here. The quantity in (2.4) is thus estimated by

$$I_n = I_n(B) = \frac{1}{n(n-1)}\sum_{j\neq i}\widehat{U}_i\widehat{U}_j\frac{1}{h}\int_B K_h\left((X_i-X_j)'\beta\right)d\beta\ .$$

Zhu and Li (1998) first proposed to use an unweighed integral of expectations conditional upon single linear indices, yielding a statistic close to, but different than, I_n for checking a linear regression model. However, they do not study the related test. Instead, their test is based on their integral statistic plus a term of the form $(1/n)\sum_{i=1}^n \widehat{U}_i\phi\left(\|X_i\|\right)$, where $\phi(\cdot)$ is the standard normal univariate density (or any other known function). Hence, they combine a test statistic based on nonparametric methods with a directional test statistic. The asymptotic behavior of their test statistic under H_0 is completely driven by the second one.

By contrast, we directly base our test on the integral statistic I_n. Let v_n^2 be the variance of $nh^{1/2}I_n$ under H_0, which is positive and finite as shown below. With at hand a consistent estimator $\widehat{v}_n^2$, an asymptotic α-level test is given by

$$\text{Reject } H_0 \quad \text{if } nh^{1/2}I_n \geq z_{1-\alpha}\widehat{v}_n\ ,$$

where $z_{1-\alpha}$ is the $(1-\alpha)$-th quantile of the standard normal distribution. The conditional variance of $nh^{1/2}I_n$ writes

$$v_n^2 = \frac{2}{n(n-1)}\sum_{j\neq i}\sigma^2(X_i)\sigma^2(X_j)h^{-1}\mathbb{E}_\beta^2\left[K_h\left((X_i-X_j)'\beta\right)\right]\ ,$$

where $\mathbb{E}_B\left[g(\beta)\right] := \int_B g(\beta)\,d\beta$ for any function $g(\cdot)$ of β. In general, the conditional variance $\sigma^2\left(\cdot\right)$ is unknown, but with at hand a nonparametric estimator such that

$$\sup_{1\leq i\leq n}\left|\frac{\widehat{\sigma}^2(X_i)}{\sigma^2(X_i)}-1\right| = o_{\mathbb{P}}(1)\ , \tag{2.7}$$

v_n^2 can be consistently estimated by

$$\widehat{v}_n^2 = \frac{2}{n(n-1)}\sum_{j\neq i}\widehat{\sigma}^2(X_i)\widehat{\sigma}^2(X_j)h^{-1}\mathbb{E}_B^2\left[K_h\left((X_i-X_j)'\beta\right)\right]\ .$$

Many nonparametric estimators could be used. For instance, one can consider

$$\widehat{\sigma}^2(x) = \frac{\sum_{i=1}^n Y_i^2 \mathbb{I}\{\|x - X_i\| \le b\}}{\sum_{i=1}^n \mathbb{I}\{\|x - X_i\| \le b\}} - \left(\frac{\sum_{i=1}^n Y_i \mathbb{I}\{\|x - X_i\| \le b\}}{\sum_{i=1}^n \mathbb{I}\{\|x - X_i\| \le b\}}\right)^2 ,$$

where b is a bandwidth parameter converging to zero as the sample size increases, which can be selected independently of h. Guerre and Lavergne (2005) then provide some primitive conditions for (2.7). It is then straightforward to show that $\widehat{v}_n^2/v_n^2 = 1 + o_{\mathbb{P}}(1)$ under H_0. Given our focus, we will proceed assuming this condition holds.

The use of a nonparametric estimator of the error's variance does not affect the test at a first order. A simpler alternative is to plug estimated parametric residuals in the expression of v_n^2 in place of the unknown variance components, which gives

$$\widehat{v}_n^2 = \frac{2}{n(n-1)} \sum_{j \ne i} \widehat{U}_i^2 \widehat{U}_j^2 h^{-1} \mathbb{E}_\beta^2 \left[K_h\left((X_i - X_j)'\beta\right)\right] .$$

This alternative estimator is consistent for v_n^2 under H_0, but overestimates it when the parametric model is incorrect, and thus likely yields some loss in power for the test. For this reason, we do not recommend its use in practice. Nevertheless, the asymptotic theory in the next section allows for its use.

2.3 Asymptotic analysis

To avoid some technicalities, the parametric regression is taken to be linear in variables. Our results can be extended to a general parametric regression, see for instance Lavergne and Patilea (2006) for assumptions and technicalities. However, we do not restrict the data to exhibit normality or homoscedasticity. We first state our general assumptions on the data-generating process, the kernel and smoothing parameter.

Assumption D *(a) The random vectors $(\varepsilon_1, X_1')', \ldots, (\varepsilon_n, X_n')'$ are independent copies of the random vector $(\varepsilon, X')' \in \mathbb{R}^{1+q}$, where $\mathbb{E}(\varepsilon \mid X) = 0$ and $\mathbb{E}(\varepsilon^4) < \infty$.*
(b) Let $\sigma^2(x) = \mathbb{E}(\varepsilon^2 \mid X = x)$. There exist constants $\underline{\sigma}^2$ and $\overline{\sigma}^2$ such that for any x $0 < \underline{\sigma}^2 \le \sigma^2(x) \le \overline{\sigma}^2 < \infty$.
(c) X is continuous with bounded density $f(\cdot)$, and the density $f_\beta(\cdot)$ of $X'\beta$ is such that for some C, $|f_\beta(\cdot)| \le C$ for any $\beta \in B$. If B is not included in a half-hypershpere, X is assumed to be bounded.
(d) Let $Z = [Z_i, i = 1, \ldots n] = [(1, X_i'), i = 1, \ldots n]$ be the design matrix. There exists a positive definite matrix A such that $n^{-1}Z'Z \xrightarrow{p} A$. $\theta \in \Theta$, a compact of $\mathbb{R}^{1+q}$.

Assumption K *(a) The kernel $K(\cdot)$ is a bounded symmetric density with $K(0) > 0$ and an integrable Fourier transform. (b) $h \to 0$ and $(nh^2)^{\alpha} / \ln n \to \infty$ for some $\alpha \in (0,1)$.*

Assumptions D(c) comes from our Lemma 2.1. It requires in particular lack of multicollinearity among the regressors. For a bounded X, a bounded density for X implies that $f_\beta(\cdot)$ is bounded uniformly in $\beta \in \mathbb{S}^q$. The assumptions on the kernel $K(\cdot)$ are satisfied by most kernels used in practice. The restrictions on the bandwidth are compatible with optimal choices for regression estimation, see e.g. Härdle and Marron (1985), and for regression checks, see Guerre and Lavergne (2002). The following theorem states the asymptotic validity of our test.

Theorem 2.2 *Under Assumptions D and K and if $\widehat{v}_n^2/v_n^2 = 1 + o_{\mathbb{P}}(1)$ under H_0, the test based on I_n has asymptotic level α conditionally on the X_i.*

Let us now investigate the ability of our test to detect directional departures from the null hypothesis. Consider a real-valued function $\delta(X)$ such that

$$\mathbb{E}[(1, X')\delta(X)] = \mathbf{0} \quad \text{and} \quad 0 < \mathbb{E}[\delta^4(X)] < \infty \, . \tag{2.8}$$

The first condition ensures that $\delta(\cdot)$ is orthogonal to any linear combination of the regressors. We do not impose smoothness restrictions on the function $\delta(\cdot)$ as is frequent in this kind of analysis. We consider the sequence of local directional alternatives

$$H_{1n} : \ \mathbb{E}\left[Y|X\right] = (1, X')\,\theta_0 + r_n\delta(X), \quad n \geq 1 \, . \tag{2.9}$$

Such directional alternatives can be detected if $r_n^2 n h^{1/2} \to \infty$, where h applies to the univariate variable defined by a single linear index in X. By comparison, when one uses a regression check based on a standard "multidimensional" nonparametric estimator, $r_n^2 n h^{q/2} \to \infty$ is needed for consistency. Hence, from the theoretical point of view, the asymptotic power of our test against directional alternatives is not affected by the dimension of the regressors.

Theorem 2.3 *Under Assumptions D and K, if $\widehat{v}_n^2/v_n^2 = O_{\mathbb{P}}(1)$ and $r_n^2 n h^{1/2} \to \infty$, the test based on I_n is consistent conditionally on the X_i against the sequence of alternatives H_{1n} with $\delta(X)$ satisfying (2.8).*

2.4 Bootstrap

While the test can be implemented using asymptotic critical values for large samples, the asymptotic approximation is likely not accurate for small or moderate samples, as is the case for most regression checks. The wild bootstrap, initially proposed by Wu (1986), is thus often used to compute small sample critical values, see e.g. Härdle and Mammen (1993) and Stute and al. (1998). Here we use a generalization of this method, the smooth conditional moments bootstrap introduced by Gozalo (1997). It consists in drawing n i.i.d. random variables ω_i independent from the original sample with $\mathbb{E}\omega_i = 0$, $\mathbb{E}\omega_i^2 = 1$, and $\mathbb{E}\omega_i^4 < \infty$, and to generate bootstrap observations of Y_i as $Y_i^* = \mu(X_i, \widehat{\theta}_n) + \widehat{\sigma}(X_i)\omega_i$, $i = 1, \ldots n$. A bootstrap test statistic is built from the bootstrap sample as the original test statistic was. When this scheme is repeated many times, the bootstrap critical value $z^*_{1-\alpha,n}$ at level α is the empirical $(1-\alpha)$-th quantile of the bootstrapped test statistic. This critical value is then compared to the initial test statistic. The following theorem can be shown following the lines of Theorem 2.2's proof.

Theorem 2.4 *Under the assumptions of Theorem 2.2 and Condition (2.7), the bootstrap critical value yields a test based on I_n with asymptotic level α conditionally on the X_i.*

3 Applications

3.1 Practical considerations

A first practical issue relates to the fact that the same bandwidth is used for all directions $X'\beta$. Hence it is desirable to transform the regressors to make different linear combinations comparable. An easy way is to center and rescale the matrix of observations on X so that it has mean zero and variance identity. Alternatively, as suggested by Bierens (1982) for the ICM test, one can map each regressor onto $(0, 1)$.

Implementation of our test requires integration on the (half) hypersphere or a subset of it. To approximate the integral in practice, it is sufficient to draw a large number of points randomly distributed on the (half) hypersphere, to evaluate the function under the integral for each draw and to compute the average. A draw can be easily performed by sampling independent $z_i, i = 1, \ldots q$, distributed as $N(0, 1)$ and to define β as the vector $z/\|z\|$. By the radial symmetry of the normal distribution, this gives points uniformly distributed on the hypersphere. In some cases, it may be possible to derive the analytic form of the integral. From the previous arguments, we have that

$$\int_{\mathbb{S}^q} K(u'\beta)\, d\beta = \int_{\mathbb{R}^q} K(\frac{u'z}{\|z\|})\phi(z)dz$$

where $\phi(\cdot)$ is the q-variate standard normal density. By a suitable change of variables, this equals

$$\int_{\mathbb{R}^q} K(\|u\|\frac{z_1}{\|z\|})\phi(z)dz \ ,$$

and thus depends only depends on $\|u\|$. However, deriving the analytic formula of this function can be quite tedious, even with symbolic computation engines, while numerical approximation is quite fast and easy. Matlab codes to implement the test are available from the authors.

3.2 Simulation study

The simulation study had two main objectives. First, we wanted to determine the sensitivity of our test to the smoothing parameter h. Second, we wanted to compare the small sample power of our test to the test of Zheng (1996) and Li and Wang (1998) on the one hand, and to the tests of Bierens (1982) and Escanciano (2006) on the other hand.

Let us first present briefly the different tests we considered. Zheng's test is based on the statistic (2.6) where $h^{-1}K_h\left((X_i - X_j)'\beta\right)$ is replaced by $h^{-q}K_h\left(X_i - X_j\right)$ with a multivariate kernel. When properly normalized, this statistic has an asymptotic standard normal distribution. Li and Wang (1998) investigated application of the wild bootstrap to this test. We used the smooth conditional moment bootstrap, which is also valid as can be shown following standard arguments. Bierens' test is based on a statistic of the form

$$n\int_{\mathbb{R}^q}\left|\frac{1}{n}\sum_{i=1}^{n} U_i(\theta)\exp(iX_i'\beta)\right|^2 \phi(\beta)d\beta = \frac{1}{n}\sum_{i,j} U_i(\theta)U_j(\theta)\exp\left(-\frac{\|X_i - X_j\|^2}{2}\right) \ ,$$

where $\phi(\beta)$ is the standard normal density on $\mathbb{R}^q$, see Bierens (1982, p. 111).* This statistic thus resembles to a kernel-based one, with a kernel depending only on the norm and a fixed bandwidth. Dominguez (2004) shows that the wild bootstrap is valid and

*Escanciano (2006) also used this form of the ICM test for comparison. In a personal communication, Herman Bierens pointed out that the asymptotic theory developed by Bierens and Ploberger (1997) applies only if the measure used in integration has compact support, so that the normal distribution should be truncated at some possibly very large values. For all practical matters however, this does not matter much.

preserves admissibility of the test, consequently we used this method to obtain critical values. Finally, Escanciano's test is based on the statistic

$$\frac{1}{n^2}\sum_{i,j} U_i(\theta)U_j(\theta)\left(\frac{1}{n}\sum_{k}\int_{\mathbb{S}^q}\mathbb{I}(X_i'\beta \le X_k'\beta)\mathbb{I}(X_j'\beta \le X_k'\beta)\,d\beta\right) ,$$

and the wild bootstrap was used to obtain critical values. Computation of the statistic was performed using Escanciano's (2006) analytic results, see his Appendix B.

We consider X with dimension four and the null hypothesis

$$H_0 \ : \ \mathbb{E}(Y|X) = (1,X)'\theta_0 \quad \text{for some } \theta_0 \ .$$

We generated samples of 100 observations from independent uniformly distributed variables for each component of X. The support was chosen as $U\left[-\sqrt{3},\sqrt{3}\right]$ to get unit variance. We sampled errors from a standard normal distribution and we constructed the response variable as

$$Y_i = (1,X_i)'\theta_0 + d\ \delta(X_i) + \varepsilon_i \quad i = 1,\ldots 100 \ ,$$

with $\theta_0 = (0.5, 0.5, 0.5, -1.5)$, and different d and $\delta(\cdot)$. For each experiment, the number of replications is 4000 under the null hypothesis and 1000 under each alternative. The number of bootstrap samples is 200 for each replication and the level is 5%. We considered the following nonparametric tests: (i) Zheng's test when the index $X'\beta_0$ is considered as the only regressor, labeled as "Zheng's test Dim 1" in our figures; (ii) Zheng's test when all four regressors are taken into account, labeled as "Zheng's test Dim 4;" (iii) our test where integration is performed on a half-hypersphere, labeled as "Our test on S;" (iv) our test where integration is performed on the subset B of the hypersphere for which the first three components of β are positive, labeled as "Our test on B." This last situation corresponds to using the qualitative information that the influences of the first three components of X on Y are of the same sign. To compute the test statistics, we used a triangular kernel with support $[-1,1]$ and we selected the bandwidth as $h = b\,n^{-2/(8+q)}$, with $q = 4$ in Case (ii) and 1 in the other cases, and b varies in $\{0.5, 1, \ldots, 3\}$. The errors' conditional variance was estimated by a kernel estimator with normal kernel and bandwidth $2n^{-1/6}$. For our test, integration was carried out on a grid of 5000 points.

In our first set of simulations, $\delta(X) = 0.1 \times (X'\beta_0/\sqrt{3})^2$, where $\beta_0 = (1,2,3,-2)/\sqrt{18}$. Figure 1 is an example of residuals plots in the case where $d = 6$. It illustrates that residuals plots are not informative on whether the model is misspecified when many regressors are present. Figure 2 compares the power of Zheng's tests and our test on the sphere when $d = 6$ and the bandwidth constant b varies. The key insight is that the performances of Zheng's test in dimension 1 is quite variable depending on the bandwidth, while our test appears to be little sensitive to this parameter. Figure 3 compares the power curves of the different tests for varying d and $b = 1$. Empirical levels are well approximated by the bootstrap for our test, but not for Zheng's test in dimension 4, with a level of 3.63. Bierens and Escanciano's tests are underrejecting, with respective levels of 3.7 and 2.1. In terms of power, there is a large loss in power for Zheng's test when going from dimension one to four. In practice however, the test based on the unknown single linear index is infeasible. The second striking fact is that our test largely outperforms Zheng's test in dimension 4, as well as the other considered tests. The curve power of our dimension-reduction test is practically indistinguishable from the one of the infeasible test. Incorporating some qualitative information further improves the power of our procedure.

In our second set of simulations, we considered the hyperbolic sine alternative $\delta(X) = \sinh(X'\beta_0/\sqrt{3})$, where $\beta_0 = (1, 2, 3, -2)/\sqrt{18}$. This alternative is particularly difficult to detect, because it resembles very much a linear function. Other features of the experiments are unchanged. Figure 4 illustrates the performances of the different nonparametric tests for $d = 6$ when the bandwidth varies. As can be seen, Zheng's tests are very sensitive to the bandwidth, while our test's power is almost stable. Figure 5 is the analog of Figure 3 for hyperbolic sine alternatives. While our test is not as powerful as Zheng's infeasible test, it outperforms all the considered competing tests. Our test on B practically yields the same power as the infeasible test. Finally, for this alternative, Bierens' ICM test does better than Escanciano's and Zheng's multidimensional test.

In a third step, we considered the sine alternative $\delta(X) = 0.1 \times \sin(\pi X'\beta_0/\sqrt{3})$, where $\beta_0 = (1, 2, 3, -2)/\sqrt{18}$. This alternative is favorable to Bierens' test, which is based on the correlation between residuals and trigonometric functions. Figure 6 compares the power curves of the different tests. As expected, Bierens' test performs better than Zheng's and Escanciano's test, but still our test does better. Integrating on B only further improves the power of our test.

To understand why our test outperforms the ICM test, recall that this procedure, as well as the one by Escanciano, estimates a quantity of the form

$$\int_{\mathbb{R}^q} \left| \mathbb{E}\left[Z\psi\left(X'u \right) \right] \right|^2 \, d\mu(u) = 0 \, .$$

From a theoretical viewpoint, they are consistent against sequences of local alternatives of the form (2.9) whenever $r_n^2 n \to \infty$, because the variance of their statistic converges at rate n. Now, instead of working with a particular known $\psi(\cdot)$ at the outset, we estimate the function of $X'\beta$ that maximizes correlation with Z. Since this function needs to be nonparametrically estimated, our test statistic has a larger variance. However, it is also expected to have a higher mean under any alternative. Since the power of the test depends of both mean and variance, our test can have higher power than its competitors, as illustrated by our experiments.

To show that our conclusions are not tied to single-index alternatives, we considered the two-indexes alternative $\delta(X) = \sinh(X'\beta_1/\sqrt{3}) + \sinh(X'\beta_2/\sqrt{3})$, where $\beta_1 = (0, 2, 1, -1)/\sqrt{6}$ and $\beta_2 = (1, 0, 2, -1)/\sqrt{6}$. As a benchmark, we took Zheng's test based on the two linear indices entering the regression function, labeled as "Zheng's test in Dim 2." Figure 7 shows that our test is close to be as powerful as Zheng's infeasible test, and is more powerful than its competitors.

Appendix

For any function $g(\cdot) \in L^1(\mathbb{R}^q) \cap L^2(\mathbb{R}^q)$, its Fourier and inverse Fourier transforms are respectively defined as $\widehat{g}(t) = (2\pi)^{-q/2} \int_{\mathbb{R}^q} \exp(it'x) g(x)\, dx$ and $(2\pi)^{-q/2} \int_{\mathbb{R}}^q \exp(-it'x)\widehat{g}(t)\, dt$. In what follows, C denotes a positive constant that may vary from line to line. We first show two lemmas that are useful for proving our main results.

Lemma 3.1 *Let $\delta(\cdot)$ be any non-zero function of X on the support of X and $h \to 0$. Under Assumptions D(c) and K(a), (i) If $\mathbb{E}\delta^2(X) < \infty$, $\mathbb{E}\left\{\delta(X_1)\delta(X_2)h^{-1}\mathbb{E}_B\left[K_h\left((X_1 - X_2)'\beta\right)\right]\right\}$ has a strictly positive finite limit. (ii) If $\mathbb{E}\delta^4(X) < \infty$ and $nh \to \infty$, then $U_n - \mathbb{E}(U_n) = o_{\mathbb{P}}(1)$ where*

$$U_n = \frac{1}{n(n-1)} \sum_{j \neq i} \delta(X_i)\delta(X_j)h^{-1}\mathbb{E}_B\left[K_h\left((X_i - X_j)'\beta\right)\right].$$

Proof. (i) Denoting by $\widehat{K}(\cdot)$ the Fourier transform of $K(\cdot)$,

$$\begin{aligned}
&\mathbb{E}\left\{\delta(X_1)\delta(X_2)h^{-1}\mathbb{E}_B\left[K_h\left((X_1 - X_2)'\beta\right)\right]\right\} \\
&\quad = (2\pi)^{-1/2}\,\mathbb{E}_B\left\{\mathbb{E}\left[\delta(X_1)\delta(X_2)h^{-1}\int \exp\left(-it(X_1 - X_2)'\beta/h\right)\widehat{K}(t)\, dt\right]\right\} \\
&\quad = (2\pi)^{q-1/2}\,\mathbb{E}_B\left\{\int \left|\widehat{\delta f}(t\beta)\right|^2 \widehat{K}(ht)\, dt\right\} .
\end{aligned}$$

As $\left|\widehat{K}(\cdot)\right| \leq \widehat{K}(0) = (2\pi)^{-1/2}$, Lebesgue's dominated convergence yields the limit

$$(2\pi)^{q-1/2}\int_{\mathbb{R}}\int_B \left|\widehat{\delta f}(t\beta)\right|^2 d\beta\, dt ,$$

provided it is finite. But the above quantity is bounded by

$$\int_{\mathbb{R}}\int_B \left|\widehat{\delta f}(t\beta)\right|^2 d\beta\, dt < \infty .$$

Finally, the limit is shown to be strictly positive as in the proof of Lemma 2.1.

(ii)
$$\begin{aligned}
\operatorname{Var}(U_n) &\leq \tfrac{C}{n}\operatorname{Var}\left[\delta(X_1)\delta(X_2)h^{-1}\mathbb{E}_B K_h\left((X_1 - X_2)'\beta\right)\right] \\
&\leq \tfrac{C}{nh}\mathbb{E}\left[\delta^2(X_1)\delta^2(X_2)h^{-1}\mathbb{E}_B K_h\left((X_1 - X_2)'\beta\right)\right] ,
\end{aligned}$$

and the above expectation converges to a finite limit from Part (i). ■

Let W be the matrix with generic element $\mathbb{E}_B\left[K_h\left((X_i - X_j)'\beta\right)\right]\mathbb{I}\left(i \neq j\right) / \left(hn(n-1)\right)$ and define its spectral radius as $\operatorname{Sp}(W) = \sup_{u \neq 0} \|Wu\|/\|u\|$.

Lemma 3.2 *Under Assumptions D(c) and K, (i)* $\operatorname{Sp}(W) = O_{\mathbb{P}}(n^{-1})$ *and (ii)* $n^2 h\|W\|^2$ *has a strictly positive limit, where $\|W\|$ denotes the Euclidean matrix norm.*

Proof. (i) For any $u \in \mathbb{R}^n$,

$$\begin{aligned}\|Wu\|^2 &= \sum_{i=1}^{n}\left(\sum_{j=1,j\neq i}^{n} w_{ij}\, u_j\right)^2 \leq \sum_{i=1}^{n}\left(\sum_{j=1,j\neq i}^{n} w_{ij}\right)\sum_{j=1,j\neq i}^{n} w_{ij}\, u_j^2 \\ &\leq \|u\|^2\left[\max_{1\leq i\leq n}\left(\sum_{j=1,j\neq i}^{n} w_{ij}\right)\right]^2 .\end{aligned}$$

Hence $n\mathrm{Sp}(W) \leq \max_{1\leq i\leq n}\sum_{j\neq i}\frac{1}{h(n-1)}\mathbb{E}_B K_h\left((X_i-X_j)'\beta\right)$. For all j, $|\mathbb{E}_B K_h\left((x-X_j)'\beta\right)| \leq C$ and $\mathrm{Var}\left[\mathbb{E}_B K_h\left((x-X_j)'\beta\right)\right] \leq C$. Thus the Bernstein inequality yields for any $t>0$

$$\begin{aligned}&\mathbb{P}\left[\left(\frac{(nh^2)^\alpha}{\ln n}\right)^{1/2}\max_{1\leq i\leq n}\left|\sum_{j\neq i}\frac{1}{(n-1)h}\mathbb{E}_B K_h\left((X_i-X_j)'\beta\right) - \mathbb{E}\left[\mathbb{E}_B K_h\left((X_i-X_j)'\beta\right)|X_i\right]\right| \geq t\right] \\ &\leq \sum_{1\leq i\leq n}\mathbb{E}\left[\mathbb{P}\left[\left|\frac{1}{(n-1)}\sum_{j\neq i}\mathbb{E}_B K_h\left((X_i-X_j)'\beta\right)\right.\right.\right. \\ &\qquad\qquad \left.\left.\left. - \mathbb{E}\left[\mathbb{E}_B K_h\left((X_i-X_j)'\beta\right)|X_i\right]\right| \geq th\left(\frac{\ln n}{(nh^2)^\alpha}\right)^{1/2} \mid X_i\right]\right] \\ &\leq 2n\exp\left(-\frac{t^2}{2}\frac{(nh^2)(\ln n)}{C((nh^2)^\alpha + th(nh^2)^{\alpha/2}(\ln n)^{1/2})}\right) \leq 2\exp\left[\ln n - \frac{t^2}{C'}(\ln n)(nh^2)^{1-\alpha}\right] \to 0 ,\end{aligned}$$

since $nh^2 \to \infty$ by Assumption K(b). Now

$$\mathbb{E}\left[h^{-1}\mathbb{E}_B K_h\left((X_i-X_j)'\beta\right)|X_i\right] = \int_B\int_{\mathbb{R}} K(u) f_\beta(X_i'\beta - hu)\, du\, d\beta$$

is bounded uniformly in i by Assumptions D(c) and K(a).
(ii) Write $n^2h\|W\|^2 = \frac{1}{(n-1)^2}\sum_{i\neq j} h^{-1}\mathbb{E}_B^2 K_h\left((X_i-X_j)'\beta\right)$. Hoeffding's (1963) inequality for U-statistics yields for any $\alpha \in (0,1)$

$$\begin{aligned}&\mathbb{P}\left[\left|\sum_{j\neq i}\frac{1}{n(n-1)h}\mathbb{E}_B^2 K_h\left((X_i-X_j)'\beta\right) - \mathbb{E}\left[\mathbb{E}_B^2 K_h\left((X_i-X_j)'\beta\right)\right]\right| \geq t\right] \\ &= \mathbb{P}\left[\left|\frac{1}{n(n-1)}\sum_{j\neq i}\mathbb{E}_B^2 K_h\left((X_i-X_j)'\beta\right) - \mathbb{E}\left[\mathbb{E}_B^2 K_h\left((X_i-X_j)'\beta\right)\right]\right| \geq th\right] \\ &\leq 2\exp\left(-\frac{t^2\,(nh^2)}{C}\right) \to 0 ,\end{aligned}$$

by Assumption K(b). We have

$$\begin{aligned}&\mathbb{E}\left[h^{-1}\mathbb{E}_B^2 K_h\left((X_i-X_j)'\beta\right)\right] \\ &= \mathbb{E}\left[h^{-1}\int_B K_h\left((X_i-X_j)'\beta\right)\, d\beta\int_B K_h\left((X_i-X_j)'\alpha\right)\, d\alpha\right] \\ &= (2\pi)^{q-1}h\int_{\mathbb{R}}\int_{\mathbb{R}}\int_B\int_B \widehat{K}(ht)\widehat{K}(hu)\left|\widehat{f}\,(t\beta+u\alpha)\right|^2 dt\, du\, d\beta\, d\alpha .\end{aligned}$$

By Assumption K(a),

$$
\begin{aligned}
& h\int_{\mathbb{R}}\int_{\mathbb{R}}\int_{B}\int_{B}|\widehat{K}(hu)|\left|\widehat{f}\left(t\beta+u\alpha\right)\right|^{2}\,dt\,du\,d\beta\,d\alpha \\
= & \int_{\mathbb{R}\times B}\left|\widehat{f}\left(t\beta\right)\right|^{2}dt\,d\beta\int_{B}d\alpha\int_{\mathbb{R}}|\widehat{K}(u)|\,du<\infty\,. \qquad (3.10)
\end{aligned}
$$

Now

$$
\begin{aligned}
& \left|h\int_{\mathbb{R}}\int_{\mathbb{R}}\int_{B}\int_{B}\left(\widehat{K}(ht)-\widehat{K}(0)\right)\widehat{K}(hu)\left|\widehat{f}\left(t\beta+u\alpha\right)\right|^{2}\,dt\,du\,d\beta\,d\alpha\right| \\
\leq & \ C\sup_{|ht|\leq M}\left|\widehat{K}(ht)-\widehat{K}(0)\right| \\
& \quad +2(2\pi)^{-1/2}h\int_{|t|\geq M/h}\int_{\mathbb{R}}\int_{B}\int_{B}|\widehat{K}(hu)|\left|\widehat{f}\left(t\beta+u\alpha\right)\right|^{2}\,dt\,du\,d\beta\,d\alpha\,.
\end{aligned}
$$

From the uniform continuity of $\widehat{K}(\cdot)$ and Equation (3.10), the right-hand side can be rendered arbitrarily small by choosing M small enough then letting h tend to zero. Therefore $\mathbb{E}\left[h^{-1}\mathbb{E}_B^2 K_h\left((X_i-X_j)'\beta\right)\right]$ tends to

$$
(2\pi)^{q-1}\widehat{K}(0)\int_{\mathbb{R}\times B}\left|\widehat{f}\left(t\beta\right)\right|^{2}dt\,d\beta\int_{B}d\alpha\int_{\mathbb{R}}\widehat{K}(u)\,du=(2\pi)^{q-1}K(0)\int_{B}d\alpha\int_{\mathbb{R}\times B}\left|\widehat{f}\left(t\beta\right)\right|^{2}dt\,d\beta\,,
$$

using Assumption K(a). ■

Proof of Theorem 2.2. Let $\varepsilon=(\varepsilon_1,\ldots,\varepsilon_n)'$. We have

$$
I_n=I_{0n}-2I_{1n}+I_{2n}=\varepsilon'W\varepsilon-2(\widehat{\theta}_n-\theta_0)'Z'W\varepsilon+(\widehat{\theta}_n-\theta_0)'Z'WZ(\widehat{\theta}_n-\theta_0)\,,
$$

Under Assumption D, $\widehat{\theta}_n-\theta_0=O_{\mathbb{P}}(n^{-1/2})$. Hence $I_{2n}\leq \mathrm{Sp}(W)\|Z(\widehat{\theta}_n-\theta_0)\|^2=O_{\mathbb{P}}(n^{-1})$ by Lemma 3.2(i). Let $\mathbb{E}_n$ denote the conditional expectation given the X_i, Z_k be any column of Z, $k=1,\ldots d+1$, and $\overline{Z}_k=Z_k'W$. Then Marcinkiewicz-Zygmund's and Minkowski's inequalities imply that there is some C independent of n such that

$$
\begin{aligned}
\mathbb{E}_n\left|Z_k'W\varepsilon\right| \ \leq & \ C\left\{\mathbb{E}_n^2\left|\sum_{i=1}^n\overline{Z}_{ki}^2\varepsilon_i^2\right|^{1/2}\right\}^{1/2}\leq C\left\{\sum_{i=1}^n\overline{Z}_{ki}^2\mathbb{E}_n^2|\varepsilon_i|\right\}^{1/2} \\
\leq & \ C\|Z_k'W\|\leq C\mathrm{Sp}(W)\|Z_k\|=O_{\mathbb{P}}(n^{-1/2})\,.
\end{aligned}
$$

Hence $I_{1n}=O_{\mathbb{P}}(n^{-1})$. Now from Lemma 2(i) by Guerre and Lavergne (2005), $nh^{1/2}I_{0n}/v_n$ converges to a standard normal conditionally on the X_i if $\|W\|^{-1}\mathrm{Sp}(W)=o_{\mathbb{P}}(1)$. Lemma 3.2 allows to conclude. ■

Proof of Theorem 2.3. Under H_{1n}, $U_i(\widehat{\theta}_n) = \varepsilon_i - Z_i(\widehat{\theta}_n - \theta_0) + r_n\delta(X_i)$. Letting $\delta = [\delta(X_1), \dots \delta(X_n)]'$, I_n can be decomposed as $I_{0n} - 2I_{1n} + I_{2n} - 2I_{3n} - 2I_{4n} + I_{5n}$, where $I_{3n} = r_n\delta' W Z(\widehat{\theta}_n - \theta_0)$, $I_{4n} = r_n \delta' W \varepsilon$, and $I_{5n} = r_n^2 \delta' W \delta$. By Assumption D(c) and Lemma 3.2(ii), $v_n^2 \leq \overline{\sigma}^4 n^2 h \|W\|^2 = O_{\mathbb{P}}(1)$. Hence $nh^{1/2} I_{0n} = O_{\mathbb{P}}(1)$. Because under our assumptions, $\widehat{\theta}_n - \theta_0 = O_{\mathbb{P}}(n^{-1/2})$, I_{1n} and I_{2n} are both $O_{\mathbb{P}}(n^{-1})$ as in Theorem 2.2's proof. Since $|u'Wv| \leq \|u\|\|v\|\mathrm{Sp}(W)$, $r_n^{-1} I_{3n} \leq \|\delta\|\|Z(\widehat{\theta}_n - \theta_0)\|\mathrm{Sp}(W) = O_{\mathbb{P}}(n^{-1/2})$. Also $I_{4n} = O_{\mathbb{P}}(r_n n^{-1/2})$ by the same arguments used for dealing with I_{1n}. Lemma 3.1(ii) yields $I_{5n} = r_n^2 C + o_{\mathbb{P}}(r_n^2)$ with $C > 0$. Collecting results, it follows that $nh^{1/2} I_n = nh^{1/2} r_n^2 C + o_{\mathbb{P}}(r_n^2 n h^{1/2})$. Deduce from $\widehat{v}_n^2 / v_n^2 = O_{\mathbb{P}}(1)$ and $r_n^2 n h^{1/2} \to \infty$ that $nh^{1/2} I_n / \widehat{v}_n$ diverges in probability. ■

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Figure 1: Quadratic alternative: Residuals plots

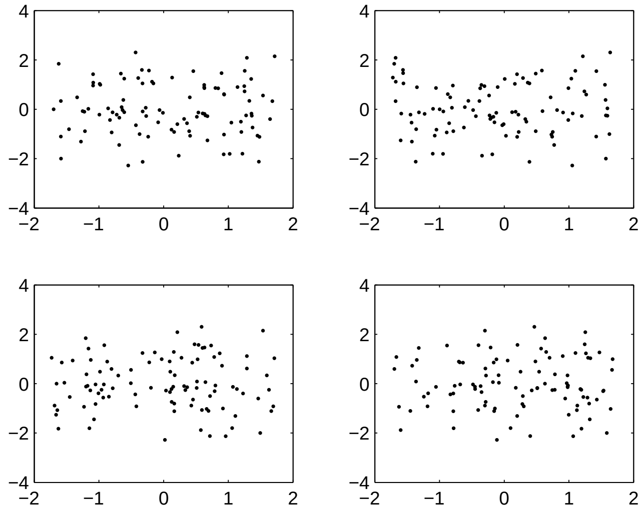

Figure 2: Quadratic alternative — varying bandwidth

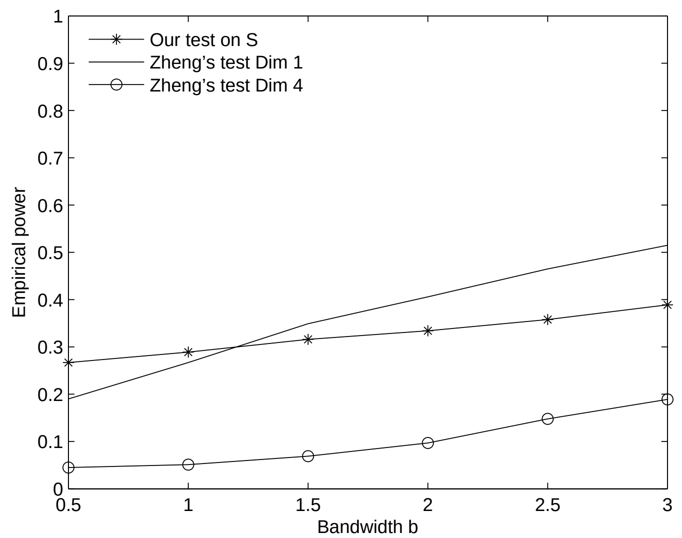

Figure 3: Quadratic varying alternative

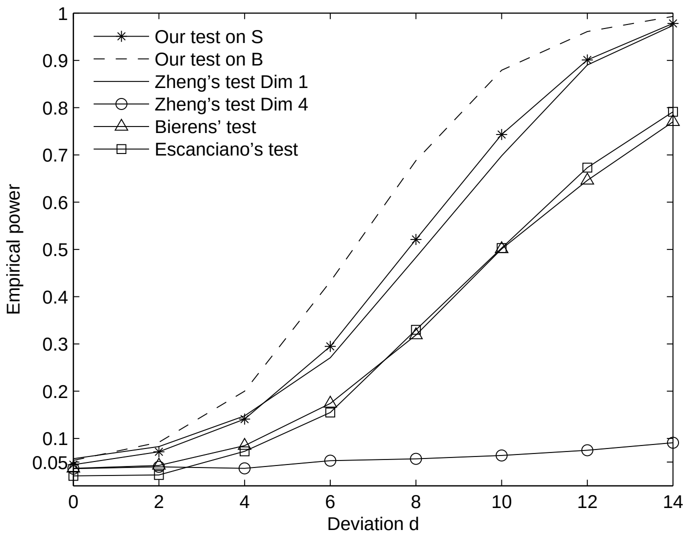

Figure 4: Sinh alternative — varying bandwidth

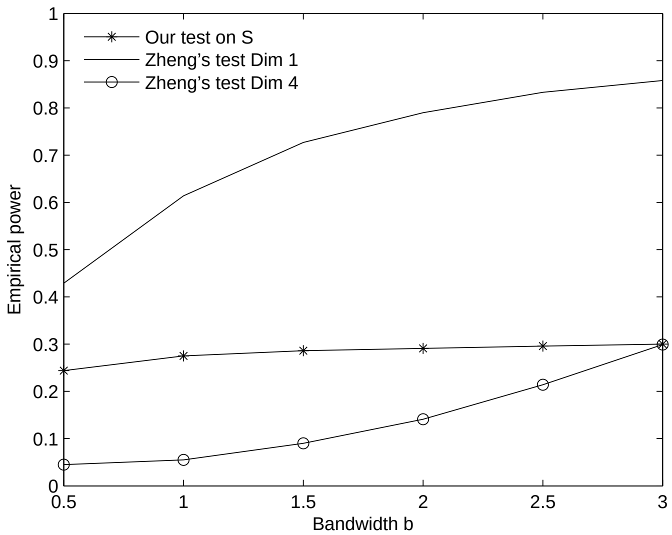

Figure 5: Sinh varying alternative

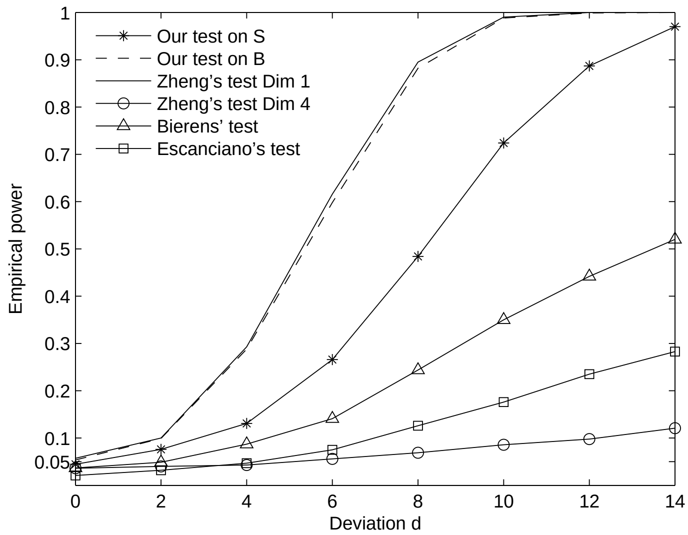

Figure 6: Sine varying alternative

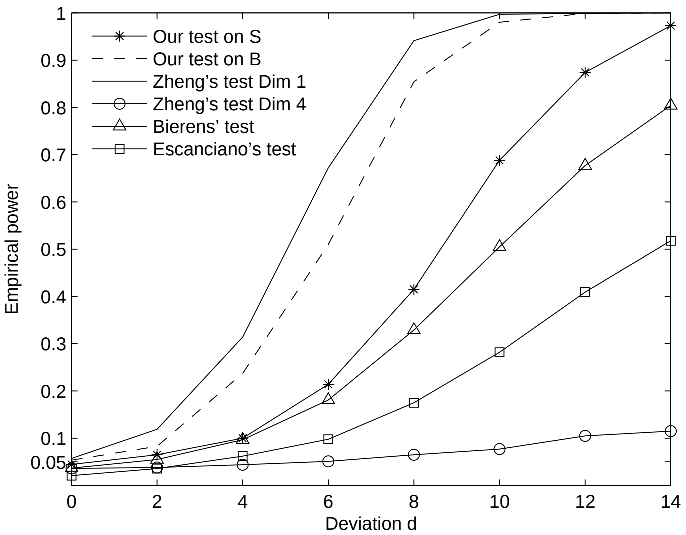

Figure 7: Two-indexes varying alternative