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SFB 649 Discussion Paper 2015-053

# Specification Testing in Random Coefficient Models

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# Specification Testing in Random Coefficient Models\*

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## Abstract

In this paper, we suggest and analyze a new class of specification tests for random coefficient models. These tests allow to assess the validity of central structural features of the model, in particular linearity in coefficients and generalizations of this notion like a known nonlinear functional relationship. They also allow to test for degeneracy of the distribution of a random coefficient, i.e., whether a coefficient is fixed or random, including whether an associated variable can be omitted altogether. Our tests are nonparametric in nature, and use sieve estimators of the characteristic function. We analyze their power against both global and local alternatives in large samples and through a Monte Carlo simulation study. Finally, we apply our framework to analyze the specification in a heterogeneous random coefficients consumer demand model.

**Keywords:** Nonparametric specification testing, random coefficients, unobserved heterogeneity, sieve minimum distance, characteristic function, consumer demand.

*JEL classification:* C12, C14

## 1 Introduction

Heterogeneity of individual agents is now widely believed to be an important - if not the most important - source of unobserved variation in a typical microeconomic application.

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Increasingly, the focus of econometrics shifts towards explicitly modeling this central feature of the model through random parameters, as opposed to searching for fixed parameters that summarize only, say, the mean effect. However, as always when additional features are being introduced, this step increases the risk of model misspecification and therefore introducing bias. This suggests to use all the information available in the data to assess the validity of the chosen specification through a test before performing the main analysis. A second important feature of a specification test is that we may be able to find a restricted model that is easier to implement than the unrestricted one. This feature is particularly important in models of complex heterogeneity, which are generically only weakly identified and therefore estimable only under great difficulties.

This paper proposes a family of nonparametric specification tests in models with complex heterogeneity. We focus on the important class of random coefficient models, i.e., models in which there is a finite ( $d_b$  dimensional) vector of continuously distributed and heterogeneous parameters  $B \in \mathbb{R}^{d_b}$ , and a known structural function  $g$  which relates these coefficients and a  $d_x$  dimensional vector of observable explanatory variables  $X$  to a continuous dependent variable  $Y$ , i.e.,

$$Y = g(X, B). \tag{1.1}$$

Throughout this paper, we assume that  $X$  is independent of  $B$  (however, as we discuss below, this does not preclude extensions where some variables in the system are endogenous). The leading example in this class of models is the linear random coefficient model, where  $g(X, B) = X'B$ , but we also propose specification tests in models where  $g$  is nonlinear. Indeed, in extensions we also consider the case where  $Y$  is binary, and/or where  $Y$  is a vector.

The simple linear model with independent random coefficients is well suited to illustrate our contribution and to explain the most important features of such a nonparametric specification test. This model is known to be exactly point identified in the sense that there is a one-to-one mapping from the conditional probability density function of the observable variables,  $f_{Y|X}$  to the density of random coefficients  $f_B$  such that the true density of random coefficients is associated with exactly one density of observables (see, e.g., Beran et al. [1996] and Hoderlein et al. [2010]). However, despite the one-to-one mapping between population density of the data and density of random coefficients, the model imposes structure that can be used to assess the validity of the model. For instance, in the very same model, the conditional expectation is linear, i.e.,  $E[Y|X] = b_0 + b_1X_1 + \dots + b_kX_k$ , where  $b_j = E[B_j]$ . This means that a standard linear model specification test for quadratic terms in  $X$ , or, somewhat more elaborate, nonparametric specification tests involving a nonparametric regression as alternative could be used to test the specification. Similarly, in this model the conditional skedastic function is at most quadratic in  $X$ , so any evidence of higher order terms can be taken as rejection of this linear random

coefficients specification, too. However, both of these tests do not use the entire distribution of the data, and hence do not allow us to discern between the truth and certain alternatives.

In contrast, our test will be based on the characteristic function of the data, i.e., we use the entire distribution of the data to assess the validity of the specification. In the example of the linear model, we compare the distance between a series least squares estimator of the unrestricted characteristic function  $E[\exp(itY)|X]$ , and an estimator of the restricted one, which is  $E[\exp(it(X'B))|X] = \int \exp(it(X'b))f_B(b)db$ , where the probability density function  $f_B$  of the random coefficients  $B$  is replaced by a sieve minimum distance estimator under the hypothesis of linearity. More specifically, using the notation  $\varepsilon(X, t) = E[\exp(itY) - \exp(it(X'B))|X]$ , our test is based on the observation that under the null hypothesis of linearity,  $\varepsilon(X, t) = 0$  holds, or equivalently,

$$\int E [|\varepsilon(X, t)|^2] \varpi(t)dt = 0,$$

for any strictly positive weighting function  $\varpi$ . Our test statistic is then given by the sample counterpart

$$S_n \equiv n^{-1} \sum_{j=1}^n \int |\hat{\varepsilon}_n(X_j, t)|^2 \varpi(t)dt,$$

where  $\hat{\varepsilon}_n$  denotes an estimator of  $\varepsilon$  as described above. We reject the null hypothesis of linearity if the statistic  $S_n$  becomes too large.

This test uses evidently the entire distribution of the data to assess the validity of the specification. It therefore implicitly uses all available comparisons between the restricted and the unrestricted model, not just the ones between, say a linear conditional mean and a nonparametric conditional mean. Moreover, it does not even require that these conditional means (or higher order moments) exist. To see that our test uses the information contained in the conditional moments, consider again the linear random coefficients model. Using a series expansion of the exponential function,  $\varepsilon(X, t) = 0$  is equivalent to

$$\sum_{l=0}^{\infty} (it)^l \{E[Y^l|X] - E[(X'B)^l|X]\} / (l!) = 0,$$

provided all moments exist. This equation holds true, if and only if, for every coefficient  $l \geq 1$  :

$$E[Y^l|X] = E[(X'B)^l|X],$$

i.e., there is equality of all of these conditional moments. This implies, in particular, the first and second conditional moment equation  $E[Y|X] = X'E[B]$  and  $E[Y^2|X] = X'E[BB']X$ . As such, our test exploits potential discrepancies in any of the conditional moments, and works even if some or all of them do not exist.

Our test is consistent against a misspecification of model (1.1) in the sense that, under the alternative, there exists no vector of random coefficients  $B$  satisfying the model equation (1.1) for a known function  $g$ . Indeed, such a misspecification leads to a deviation of the unrestricted from the restricted conditional characteristic function. Moreover, our test is also consistent against certain specific other alternatives, e.g., if the null is the linear random coefficient model and the alternative is a higher order polynomial with random parameters.

However, we can also use the same testing principle to analyze whether or not a parameter is nonrandom, which usually allows for a  $\sqrt{n}$  consistent estimator for this parameter, and whether it has in addition mean zero which implies that we may omit the respective variable altogether. This is important, because from a nonparametric identification perspective random coefficient models are weakly identified (i.e., stem from the resolution of an ill posed inverse problem), a feature that substantially complicates nonparametric estimation<sup>1</sup>. If we think of a parametric model as an approximation to a more complex nonparametric model, this is likely also going to affect the finite sample behavior of any parametric estimator. Since this can be tied to the number of dimensions (see Hoderlein et al. [2010]), it is desirable to reduce the number of dimensions of random parameters as much as possible. Our test may serve as guidance in this process.

Finally, it is important to note that our method also applies to other point identified random coefficient models such as models that are linear in parameters, but where  $X$  is replaced by a element-wise transformation of the covariates (i.e.,  $X_j$  is replaced by  $h_j(X_j)$  with unknown  $h_j$ . See Gautier and Hoderlein [2015] for the formal argument that establishes identification). The reason is that the mean regression in these models is still of an additive structure, i.e., in particular it does preclude interaction terms among the variables that feature across all moments.

**Extensions.** While setting up the basic framework is a contribution in itself, a key insight in this paper is that testing is possible even if the density of random coefficients is not point identified under the null hypothesis. This is important, because many structural models are not linear in an index. As such, it is either clear that they are not point identified in general and at best set identified (see Hoderlein et al. [2014], for such an example), or identification is unknown. To give an example of such a model, consider a stylized version of the workhorse QUARDS model of consumer demand (Banks et al. [1997]), where demand for a good  $Y$  is defined through:

$$Y = B_0 + B_1X + B_2X^2,$$

where  $B_j$  denotes parameters, and  $X$  log total expenditure. For reasons outlined in Masten

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<sup>1</sup>In a nonparametric sense, there is a stronger curse of dimensionality associated with random coefficient models than with nonparametric density estimation problems (see, e.g., Hoderlein et al. [2010]).

[2015], the joint density of random parameters  $B_0, B_1, B_2$  is not point identified in general. Our strategy is now to solve a functional minimization problem that minimizes a similar distance as outlined above between restricted and unrestricted model, and allows us to obtain one element in this set as minimizer. If the distance between the restricted model and the unrestricted model is larger than zero, we conclude that we can reject the null that the model is, in our example, a heterogeneous QUAIDS. However, if the distance is not significantly different from zero, there still may be other non-QUAIDS models which achieve zero distance, and which we therefore cannot distinguish from the heterogeneous QUAIDS model. As such, in the partially identified case we do not have power against all possible alternatives, and our test becomes conservative.

Interestingly, even if our model is not identified under the null hypothesis, such as in the case of the random coefficients QUAIDS model, our test still has power against certain alternatives, e.g., any higher polynomial random coefficient model. Again, since our test compares all conditional moments,  $\varepsilon(X, t) = 0$  for all  $t$  implies that the cubic model  $Y = \tilde{B}_0 + X\tilde{B}_1 + X^2\tilde{B}_2 + X^3\tilde{B}_3$  with random coefficients  $(\tilde{B}_0, \tilde{B}_1, \tilde{B}_2, \tilde{B}_3)$  or any other higher polynomial model is misspecified. In this sense, our test has power even in situations where neither the null nor the alternative model is identified.

The second extension is that our testing principle extends to systems of equations, i.e., situations in which the endogenous variable is not a scalar, but a vector, by replacing the scalar conditional characteristic function with a vector valued one, i.e.,  $E[\exp(it'Y)|X = x]$ . For instance, one may reformulate the triangular random coefficients model of Hoderlein et al. [2014], where  $Y_1 = A_0 + A_1Y_2$ ,  $Y_2 = B_2 + B_3X$  as

$$\begin{aligned} Y_1 &= B_0 + B_1X, \\ Y_2 &= B_2 + B_3X, \end{aligned}$$

with  $B = (B_0, B_1, B_2, B_3) \perp X$ , and then either use the minimum distance principle outlined above, or, under the additional assumptions in Hoderlein et al. [2014], their estimator for the restricted model.

Finally, we may extend the approach outlined in this paper to binary or discrete dependent variables, provided we have a special regressor  $Z$ , as in Lewbel [2000]. In this case, we replace the density of the data with the marginal probability with respect to the special regressor; otherwise, most of the above reasoning remains virtually unchanged.

**Related Literature.** As already mentioned, this paper draws upon several literatures. The first is nonparametric random coefficients models, a recently quite active line of work, including work on the linear model (Beran and Hall [1992], Beran et al. [1996], and Hoderlein et al. [2010]),

the binary choice model (Ichimura and Thompson [1998] and Gautier and Kitamura [2013]), and the treatment effects model (Gautier and Hoderlein [2015]). Related is also the wider class of models analyzed in Fox and Gandhi [2009] and Lewbel and Pendakur [2013], who both analyze nonlinear random coefficient models, Masten [2015] and Matzkin [2012], who both discuss identification of random coefficients in a simultaneous equation model, Hoderlein et al. [2014] who analyze a triangular random coefficients model, and Dunker et al. [2013] and Fox and Lazzati [2012] who analyze games.

As far as we know, the general type of specification tests we propose in this paper is new to the literature. In linear semiparametric random coefficient models, Beran [1993] proposes a minimum distance estimator for the unknown distributional parameter of the random coefficient distribution. Within this framework of a parametric joint random coefficients' distribution, Beran also proposes goodness of fit testing procedures. Also, in a parametric setup where the unknown random coefficient distribution follows a parametric model, Swamy [1970] establishes a test for equivalence of random coefficient across individuals, i.e., a test for degeneracy of the random coefficient vector. We emphasize that with our testing methodology, despite less restrictive distributional assumptions, we are able to test degeneracy of a subvector of  $B$  while others are kept as random. Another test in linear parametric random coefficient models was proposed by Andrews [2001], namely a test for degeneracy of some random coefficients. In contrast, our nonparametric testing procedure is based on detecting differences in conditional characteristic function representation and, as we illustrate below, we do not obtain boundary problems as in Andrews [2001].

In this paper, we use sieve estimators for the unknown distributional elements. In the econometrics literature, sieve methodology was recently used to construct Wald statistics (see Chen and Pouzo [2015] and Chen and Pouzo [2012] for sieve minimum distance estimation) or nonparametric specification tests (see Breunig [2015b]), and, in nonparametric instrumental regression, tests based on series estimators have been proposed by Horowitz [2012] and Breunig [2015a]. Moreover, in the nonparametric IV model, tests for parametric specification have been proposed by Horowitz [2006] and Horowitz and Lee [2009], while Blundell and Horowitz [2007] proposes a test of exogeneity. Santos [2012] develops hypothesis tests which are robust to a failure of identification. More generally, there is a large literature on model specification tests based on nonparametric regression estimators in  $L^2$  distance starting with Härdle and Mammen [1993]. Specification tests in nonseparable were proposed by Hoderlein et al. [2011] and Lewbel et al. [2015]. None of these tests is applicable to specification testing in random coefficient models.

Finally, our motivation is partly driven by consumer demand, where heterogeneity plays an important role. Other than the large body of work reviewed above we would like to mention



the recent work by Hausman and Newey [2013], Blundell et al. [2010], see Lewbel [1999] for a review of earlier work.

**Overview of Paper.** In the second section, we introduce our test formally, and discuss its large sample properties in the baseline scenario. We distinguish between general specification tests, and subcases where we can separate additively a part of the model which contains only covariates and fixed coefficients from the remainder. In the third section, we focus on the extensions discussed above. The finite sample behavior is investigated through a Monte Carlo study in the fourth section. Finally, we apply all concepts to analyze the validity of a heterogeneous QUAIDS (Banks et al. [1997]) model which is the leading parametric specification in consumer demand.

## 2 The Test Statistic and its Asymptotic Properties

### 2.1 Examples of Testable Hypotheses

In the wider class of models encompassed by (1.1), we consider two different types of hypotheses tests. First, we provide a general test for the hypothesis that the structural relation between the covariates, the random coefficients, and the outcome variable coincides with a known function  $g$ . We thus consider the hypothesis <sup>2</sup>

$$H_{\text{mod}} : Y = g(X, B) \quad \text{for some random parameters } B.$$

The alternative hypothesis is given by  $P(Y \neq g(X, B) \quad \text{for all random parameters } B) > 0$ . An important example is testing the hypothesis of linearity, i.e., whether with probability one

$$H_{\text{lin}} : Y = X'B,$$

in which case the distribution of  $B$  is point identified. Another example is a quadratic form of the function  $g$  in each component of the vector of covariates  $X$ , i.e., we want to assess the null hypothesis

$$H_{\text{quad}} : Y = B_0 + X'B_1 + (X^2)'B_2,$$

for some  $B = (B_0, B_1, B_2)$ , where the square of the vector  $X$  is understood element-wise. Note that in the latter example the distribution of the random parameters  $B$  is only partially identified. As already discussed above, this fact will generally result in a lack of power against certain alternatives.

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<sup>2</sup>Equalities involving random variables are understood as equalities with probability one, even if we do not say so explicitly.

The second type of hypotheses our test allows to consider is whether a subvector of  $B$ , say,  $B_2$ , is deterministic (or, equivalently, has a degenerate distribution). More specifically, we want to consider the following hypothesis

$$H_{\text{deg}} : B_2 = b_2 \quad \text{for some } B = (B_0, B_1, B_2) \text{ satisfying (1.1).}$$

The alternative is given by  $P(B_2 \neq b_2 \text{ for all } B \text{ satisfying (1.1)}) > 0$ . While the hypothesis  $H_{\text{deg}}$  could be considered in more general models, motivated by the linear (or polynomial) model we will confine ourselves to functions  $g$  that are additively separable in the sense that

$$H_{\text{add}} : Y = g_1(X, B_{-2}) + g_2(X, B_2), \tag{2.1}$$

where  $g_1$  and  $g_2$  denote two known functions, and we use the notation  $B_{-2} = (B_0, B_1)'$ . The leading example for this type of hypothesis is of course when  $g_1$  is a linear function of a subvector  $X_1$  of covariates  $X$ , in which case we obtain a partially linear structure, i.e.,

$$H_{\text{part-lin}} : Y = B_0 + X_1' B_1 + g_2(X, B_2), \tag{2.2}$$

where  $g_2$  is a known function. This covers the following examples of hypotheses already outlined in the introduction: First, in a linear model, i.e.,  $Y = B_0 + X_1' B_1 + X_2' B_2$ , it allows to test whether the coefficient on  $X_2$  is deterministic, i.e., we may test the null

$$H_{\text{deg-lin}} : Y = B_0 + X_1' B_1 + X_2' b_2,$$

against the alternative that  $B_2$  is random. Obviously, in this case  $b_2$  is identified by standard linear mean regression identification conditions. A second example arises if, in the quadratic model, we want to test a specification with deterministic second order terms, i.e.

$$H_{\text{deg-quad}} : Y = B_0 + X_1' B_1 + (X_1^2)' b_2,$$

against the alternative that  $B_2$  is random. Note that in the latter two hypotheses, identification of  $b_2$  follows as in parametric mean regression, and in equation (2.2), point identification under the null holds for instance if  $g_2(X, b_2) = h(X_2)' b_2$  for some vector valued function  $h$  such that the associated rank condition is satisfied. In the Monte Carlo study and the application, we will only consider the case where  $b_2$  is point identified, which we consider to be the leading case. However, we would like to point out that the test applies also more generally to situations where  $b_2$  need not to be point identified, as is in the most general case defined by hypothesis  $H_{\text{add}}$ , albeit with a loss of power against some alternatives.

## 2.2 The Test Statistic

Our test statistic is based on the  $L^2$  distance between an unrestricted conditional characteristic function and a restricted one. We show below that each null hypothesis is then equivalent to

$$\varepsilon(X, t) = 0 \text{ for all } t, \quad (2.3)$$

where  $\varepsilon : \mathbb{R}^{d_x+1} \rightarrow \mathbb{C}$  is a complex valued, measurable function. Our testing procedure is based on the  $L^2$  distance of  $\varepsilon$  to zero. Equation (2.3) is equivalent to

$$\int E[|\varepsilon(X, t)|^2] \varpi(t) dt = 0,$$

for some strictly positive weighting function  $\varpi$  with  $\int \varpi(t) dt < \infty$ . In the following examples, we provide explicit forms for the function  $\varepsilon$ . The analysis is based on the assumption of independence of covariates  $X$  and random coefficients  $B$ . See also the discussion after Assumption 1 below.

**Example 1** (Testing functional form restrictions). *The null hypothesis  $H_{mod}$  is equivalent to the following equation involving conditional characteristic functions*

$$E[\exp(itY)|X] = \int \exp(itg(X, b)) f_B(b) db,$$

for each  $t \in \mathbb{R}$ , a known function  $g$ , and some random parameters  $B$ , with probability density function (p.d.f.)  $f_B$ . Hence, equation (2.3) holds true with

$$\varepsilon(X, t) = E[\exp(itY)|X] - \int \exp(itg(X, b)) f_B(b) db. \quad (2.4)$$

As already mentioned, if the function  $g$  is nonlinear the p.d.f. of random coefficients  $B$  is not necessarily point identified. On the other side, if  $g$  is the inner product of its entries, then (2.3) holds true with

$$\varepsilon(X, t) = E[\exp(itY)|X] - \int \exp(itX'b) f_B(b) db,$$

and in this case the distribution of  $B$  is point identified (see, e.g., Hoderlein et al. [2010]).

While our test, based on the function  $\varepsilon$ , is in general consistent against a failure of the null hypothesis  $H_{mod}$ , it is also consistent against certain alternative models such as higher order polynomials which are not point identified. To illustrate this, consider testing linearity of the random coefficient QUAIDS model which is given by  $Y = \tilde{B}_0 + \tilde{B}_1 X + \tilde{B}_2 X^2$  for random coefficients  $\tilde{B}_0$ ,  $\tilde{B}_1$ , and  $\tilde{B}_2$  (also independent of  $X$ ). In this case, the conditional first and second moment equation implied by equation (2.3) yield  $E[\tilde{B}_2] = 0$  and  $Var(\tilde{B}_2) = 0$ , respectively. We

thus conclude that  $\tilde{B}_2 = 0$  with probability one.

**Example 2** (Testing degeneracy under the random coefficients specification). *In the case of an additively separable structure  $H_{add}$  (see equation (2.1)), the null hypothesis  $H_{deg}$  implies the equality of conditional characteristic functions, i.e.,*

$$E[\exp(itY)|X] = \int \exp(itg_1(X, b_{-2}))f_{B_{-2}}(b_{-2})db_{-2} \exp(itg_2(X, b_2)), \quad (2.5)$$

for each  $t \in \mathbb{R}$ . Therefore, equation (2.3) holds with

$$\varepsilon(X, t) = E[\exp(itY)|X] - \int \exp(itg_1(X, b_{-2}))f_{B_{-2}}(b_{-2})db_{-2} \exp(itg_2(X, b_2)).$$

Given a partially linear structure  $H_{part-lin}$  (see equation (2.2)), the null hypothesis  $H_{deg}$  implies the equality of conditional characteristic functions, i.e., equation (2.3) holds with

$$\varepsilon(X, t) = E[\exp(itY)|X] - \int \exp(itX'_1 b_{-2})f_{B_{-2}}(b_{-2})db_{-2} \exp(itg_2(X, b_2)),$$

where the distribution of the random coefficients is identified. Our test, based on the function  $\varepsilon$ , has power against any failure of hypothesis  $H_{deg}$  if the distribution of the random coefficients under the maintained hypothesis  $H_{add}$  is identified, i.e., if  $g_1$  and  $g_2$  are linear in  $X_1$  and  $X_2$ , respectively, or element-wise transformations of each component of these vectors (see Gautier and Hoderlein [2015]).

To illustrate that our test of degeneracy has power in the random coefficient QUAIDS model  $Y = \tilde{B}_0 + \tilde{B}_1 X + \tilde{B}_2 X^2$ , note that under the null the conditional first and second moment regressions implied by equation (2.3) already yield that  $E[\tilde{B}_2] = b_2$  and  $E[\tilde{B}_2^2] = b_2^2$ , respectively. From this observation we are already in the position to conclude that  $\tilde{B}_2$  is degenerate with  $\tilde{B}_2 = b_2$ .

**Example 3** (Testing degeneracy under additive separability alone). *We also present an alternative test of degeneracy under  $H_{add}$  (see equation (2.1)) when  $g_1$  depends on covariates  $X_1$  but not on a subvector  $X_2$  of the covariates  $X = (X'_1, X'_2)'$ . In this case, we rely on additive separability alone and base our test on*

$$E[\exp(itY)|X] = E[\exp(it(Y - g_2(X, b_2))|X_1] \exp(itg_2(X, b_2)). \quad (2.6)$$

Of course, such a test is only reasonable if the sigma algebra generated by  $X$  is not contained in the one generated by  $X_1$ . This rules out, for instance, testing degeneracy in the random coefficient QUAIDS model where  $X$  is scalar and  $g_2$  is a quadratic function of  $X$ .

This test would not require any structure on the first term (despite not depending on  $X_2$ ), i.e., in equation (2.1) we do neither have to know  $g_1$ , nor would have to assume that  $B_{-2}$  is finite. In contrast to the setting in Example 2, however, we require  $b_2$  to be point identified, which in the absence of any structure on  $g_1$  may be difficult to establish. There are examples where this structure could be useful. Consider for instance a model which has a complex nonlinear function in  $X_1$ , but is linear in  $X_2$ , i.e.,  $Y = g_1(X_1, B_{-2}) + X_2' B_2$ , with an unknown function  $g_1$ . Suppose a researcher wants to test the null that the random coefficients  $B_2$  has a degenerate distribution. In this case,  $b_2$  is identified by a partially linear mean regression model, since  $E[Y|X] = \mu(X_1) + X_2' b_2$ , where  $\mu(X_1) = E[g_1(X_1, B_{-2})|X_1]$ . Evidently, this test requires less structure on the way  $X_1$  enters, but in return suffers from lower power, e.g., if  $X_1$  indeed enters through a random coefficients specification.

As already mentioned, we use the fact that equation (2.3) is equivalent to

$$\int E[|\varepsilon(X, t)|^2] \varpi(t) dt = 0,$$

for some strictly positive weighting function  $\varpi$ . Our test statistic is given by the sample counterpart to this expression, which is

$$S_n \equiv n^{-1} \sum_{j=1}^n \int |\hat{\varepsilon}_n(X_j, t)|^2 \varpi(t) dt,$$

where  $\hat{\varepsilon}_n$  is a consistent estimator of  $\varepsilon$ . Below, we show that the statistic  $S_n$  is asymptotically normally distributed after standardization. As the test is one sided, we reject the null hypothesis at level  $\alpha$  when the standardized version of  $S_n$  is larger than the  $(1 - \alpha)$ -quantile of  $\mathcal{N}(0, 1)$ .

We consider a series estimator for the conditional characteristic function of  $Y$  given  $X$ , i.e.,  $\varphi(x, t) \equiv E[\exp(itY)|X = x]$ . To do so, let us introduce a vector of basis functions denoted by  $p_m(\cdot) = (p_1(\cdot), \dots, p_m(\cdot))'$  for some integer  $m \geq 1$ . Further, let  $\mathbf{X}_m \equiv (p_m(X_1), \dots, p_m(X_n))'$  and  $\mathbf{Y}_n(t) = (\exp(itY_1), \dots, \exp(itY_n))$ . We replace  $\varphi$  by the series least squares estimator

$$\hat{\varphi}_n(x, t) \equiv p_{m_n}(x) (\mathbf{X}_{m_n}' \mathbf{X}_{m_n})^{-1} \mathbf{X}_{m_n}' \mathbf{Y}_n(t),$$

where the integer  $m_n$  increases with sample size  $n$ . We compare this unrestricted conditional expectation estimator to a restricted one which depends on the hypothesis under consideration.

**Example 4** (Testing functional form restrictions (cont.)). *Let us introduce the integral transform  $(\mathcal{F}_g f)(X, t) \equiv \int \exp(itg(X, b)) f(b) db$ , which coincides with the Fourier transform evaluated at  $tX$ , if  $g$  is linear.<sup>3</sup> If  $g$  is nonlinear, then the random coefficient's p.d.f.  $f_B$  does not*

<sup>3</sup>The Fourier transform is given by  $(\mathcal{F}\phi)(t) \equiv \int \exp(itz)\phi(z)dz$  for a function  $\phi \in L^1(\mathbb{R}^d)$  while its inverse is

need to be identified through  $\varphi = \mathcal{F}_g f$ . We estimate the function  $\varepsilon$  by

$$\widehat{\varepsilon}_n(X_j, t) = \widehat{\varphi}_n(X_j, t) - (\mathcal{F}_g \widehat{f}_{B_n})(X_j, t),$$

where the estimator  $\widehat{f}_{B_n}$  is a sieve minimum distance estimator given by

$$\widehat{f}_{B_n} \in \arg \min_{f \in \mathcal{B}_n} \left\{ \sum_{j=1}^n \int |\widehat{\varphi}_n(X_j, t) - (\mathcal{F}_g f)(X_j, t)|^2 \varpi(t) dt \right\} \quad (2.7)$$

and  $\mathcal{B}_n = \{\phi(\cdot) = \sum_{l=1}^{k_n} \beta_l q_l(\cdot)\}$  is a linear sieve space of dimension  $k_n < \infty$  with basis functions  $\{q_l\}_{l \geq 1}$ . Here,  $k_n$  and  $m_n$  increase with sample size  $n$ . As we see below, we require that  $m_n$  increases faster than  $k_n$ . Next, using the notation  $\mathbf{F}_n(t) = ((\mathcal{F}_g q_{k_n})(X_1, t), \dots, (\mathcal{F}_g q_{k_n})(X_n, t))'$ , the minimum norm estimator of  $f_B$  given in (2.7) coincides with  $\widehat{f}_{B_n}(\cdot) = q_{k_n}(\cdot)' \widehat{\beta}_n$  where

$$\widehat{\beta}_n = \left( \int \mathbf{F}_n(t)' \mathbf{F}_n(t) \varpi(t) dt \right)^{-} \int \mathbf{F}_n(t)' \Phi_n(t) \varpi(t) dt$$

and  $\Phi_n(t) = (\widehat{\varphi}_n(X_1, t), \dots, \widehat{\varphi}_n(X_n, t))'$ .<sup>4</sup> The exponent  $-$  denotes the Moore–Penrose generalized inverse. As a byproduct, we thus extend the minimum distance estimation principle of Beran and Millar [1994] to nonlinear random coefficient models and the sieve methodology.

**Example 5** (Testing degeneracy under the random coefficients specification (cont.)). We estimate the function  $\varepsilon$  by

$$\widehat{\varepsilon}_n(X_j, t) = \widehat{\varphi}_n(X_j, t) - (\mathcal{F}_{g_1} \widehat{f}_{B_{-2,n}})(X_j, t) \exp(itg_2(X_j, \widehat{b}_{2n})),$$

where the estimators  $\widehat{f}_{B_{-2,n}}$  and  $\widehat{b}_{2n}$  are a sieve minimum distance estimators of the p.d.f.  $f_{B_{-2}}$  and the parameter  $b_2$ , respectively, given by

$$(\widehat{f}_{B_{-2,n}}, \widehat{b}_{2n}) \in \arg \min_{(f,b) \in \mathcal{B}_{-2,n} \times \mathcal{B}_2} \left\{ \sum_{j=1}^n \int |\widehat{\varphi}_n(X_j, t) - (\mathcal{F}_{g_1} f)(X_j, t) \exp(itg_2(X_j, b))|^2 \varpi(t) dt \right\} \quad (2.8)$$

and  $\mathcal{B}_{-2,n} = \{\phi(\cdot) = \sum_{l=1}^{k_n} \beta_l q_l(\cdot)\}$  is a linear sieve space of dimension  $k_n < \infty$  with basis functions  $\{q_l\}_{l \geq 1}$  of  $B_{-2}$  and  $\mathcal{B}_2$  is a compact parameter space. See also Ai and Chen [2003] for sieve minimum distance estimation for finite dimensional parameters and nonparametric functions. As in the previous example,  $k_n$  and  $m_n$  increase with sample size  $n$ , but we require that  $m_n$  increases faster than  $k_n$ .

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$(\mathcal{F}^{-1} \widetilde{\phi})(z) \equiv (2\pi)^{-d} \int \exp(-itz) \widetilde{\phi}(t) dt.$

<sup>4</sup>The integral transform  $\mathcal{F}_g$  of a vector of functions is always understood element-wise, i.e.,  $(\mathcal{F}_g q_{k_n})(X_j, t) = ((\mathcal{F}_g q_1)(X_j, t), \dots, (\mathcal{F}_g q_{k_n})(X_j, t))'$ .

**Example 6** (Testing degeneracy under additive separability alone (cont.)). Let  $\widehat{b}_{2n}$  denote a consistent estimator of the point identified parameter  $b_2$ . For instance, under the partially linear structure  $H_{part-lin}$  (see equation (2.2)), we have the moment restriction  $E[Y|X] = b_0 + X_1' b_1 + g_2(X, b_2)$  and thus,  $\widehat{b}_{2n}$  would coincide with the nonlinear least squares estimator of  $b_2$ . We denote  $p_{\underline{k}_n}(\cdot) = (p_1(\cdot), \dots, p_{k_n}(\cdot))'$  and  $\mathbf{X}_{1n} \equiv (p_{\underline{k}_n}(X_{11}), \dots, p_{\underline{k}_n}(X_{1n}))'$  which is a  $n \times k_n$  matrix. Consequently, we estimate the function  $\varepsilon$  by

$$\widehat{\varepsilon}_n(X_j, t) = \widehat{\varphi}_n(X_j, t) - p_{\underline{k}_n}(X_{1j})' (\mathbf{X}_{1n}' \mathbf{X}_{1n})^{-1} \mathbf{X}_{1n}' \mathbf{U}_n \exp(itg_2(X_j, \widehat{b}_{2n})),$$

where  $\mathbf{U}_n = (\exp(it(Y_1 - g_2(X_1, \widehat{b}_{2n}))), \dots, \exp(it(Y_n - g_2(X_n, \widehat{b}_{2n}))))'$ .

## 2.3 The Asymptotic Distribution of the Statistic under the Null Hypothesis

As a consequence of the previous considerations, we distinguish between two main hypotheses, i.e., functional form restrictions and degeneracy of some random coefficients. Both types of tests require certain common assumptions, and we start out this section with a subsection where we discuss the assumptions we require in both cases. Thereafter, we analyze each of the two types of tests in a separate subsection, and provide additional assumptions to obtain the test's asymptotic distribution under each null hypothesis.

### 2.3.1 General Assumptions for Inference

**Assumption 1.** *The random vector  $X$  is independent of  $B$ .*

Assumption 1 is crucial for the construction of our test statistic. Full independence is commonly assumed in the random coefficients literature (see, for instance, Beran [1993], Beran et al. [1996], Hoderlein et al. [2010], or any of the random coefficient references mentioned in the introduction). It is worth noting that this assumption can be relaxed by assuming independence of  $X$  and  $B$  conditional on additional variables that are available to the econometrician, allowing for instance for a control function solution to endogeneity as in Hoderlein and Sherman [2015], or simply controlling for observables in the spirit of the unconfoundedness assumption in the treatment effects literature. Further,  $\mathcal{X}$  denotes the support of  $X$ .

**Assumption 2.** (i) *We observe a sample  $((Y_1, X_1), \dots, (Y_n, X_n))$  of independent and identically distributed (i.i.d.) copies of  $(Y, X)$ . (ii) *There exists a strictly positive and nonincreasing sequence  $(\lambda_n)_{n \geq 1}$  such that, uniformly in  $n$ , the smallest eigenvalue of  $\lambda_n^{-1} E[p_{m_n}(X)p_{m_n}(X)']$  is bounded away from zero. (iii) *There exists a constant  $C \geq 1$  and a sequence of positive integers  $(m_n)_{n \geq 1}$  satisfying  $\sup_{x \in \mathcal{X}} \|p_{\underline{m}_n}(x)\|^2 \leq C m_n$  with  $m_n^2 \log n = o(n\lambda_n)$ .***

Assumption 2 (ii) – (iii) restricts the magnitude of the approximating functions  $\{p_l\}_{l \geq 1}$  and imposes nonsingularity of their second moment matrix. Assumption 2 (iii) holds, for instance, for polynomial splines, Fourier series and wavelet bases. Moreover, this assumption ensures that the smallest eigenvalue of  $E[p_{\underline{m}_n}(X)p_{\underline{m}_n}(X)']$  is not too small relative to the dimension  $m_n$ . In Assumption 2 (ii), we assume that the eigenvalues of the matrix  $E[p_{\underline{m}_n}(X)p_{\underline{m}_n}(X)']$  may tend to zero at the rate  $\lambda_n$  which was recently also assumed by Chen and Christensen [2015]. On the other hand, the sequence  $(\lambda_n)_{n \geq 1}$  is bounded away from zero if  $\{p_l\}_{l \geq 1}$  forms an orthonormal basis on the compact support of  $X$  and the p.d.f. of  $X$  is bounded away from zero (cf. Proposition 2.1 of Belloni et al. [2015]). The next result provides sufficient condition for Assumption 2 (ii) to hold even if the sequence of eigenvalues  $(\lambda_n)_{n \geq 1}$  tends to zero.

**Proposition 1.** *Assume that  $\{p_l\}_{l \geq 1}$  forms an orthonormal basis on  $\mathcal{X}$  with respect to a measure  $\nu$ . Let  $(\lambda_n)_{n \geq 1}$  be a sequence that tends to zero. Suppose that, for some constant  $0 < c < 1$ , for all  $n \geq 1$  and any vector  $a_n \in \mathbb{R}^{m_n}$  the inequality*

$$\int (a_n' p_{\underline{m}_n}(x))^2 1\{f(x) < \lambda_n\} \nu(dx) \leq c \int (a_n' p_{\underline{m}_n}(x))^2 \nu(dx) \quad (2.9)$$

holds, where  $f = dF_X/d\nu$ . Then, Assumption 2 (ii) is satisfied.

Condition (2.9) is violated, for instance, if  $dF_X/d\nu$  vanishes on some subset  $\mathcal{A}$  of the support of  $\nu$  with  $\nu(\mathcal{A}) > 0$ . Estimation of conditional expectations with respect to  $X$  is more difficult when the marginal p.d.f.  $f_X$  is close to zero on the support  $\mathcal{X}$ . In this case, the rate of convergence will slow down relative to  $\lambda_n$  (see Lemma 2.4 in Chen and Christensen [2015] in case of series estimation). As we see from Proposition 2.9,  $\lambda_n$  plays the role of a truncation parameter used in kernel estimation of conditional densities to ensure that the denominator is bounded away from zero.

To derive our test's asymptotic distribution, we standardize  $S_n$  by subtracting the mean and dividing through a variance which we introduce in the following. Let  $V \equiv (Y, X)$ , and denote by  $\delta$  a complex valued function which is the difference of  $\exp(itY)$  and the restricted conditional characteristic function, i.e.,  $\delta(V, t) = \exp(itY) - (\mathcal{F}_g f_B)(X, t)$  in case of  $H_{\text{mod}}$ , and  $\delta(V, t) = \exp(itY) - E[\exp(it(B_0 + X_1' B_1)) | X_1] \exp(itg_2(X, b_2))$  in case of  $H_{\text{deg}}$ . Moreover, note that  $\int E[\delta(V, t) | X] \varpi(t) dt = 0$  holds.

**Definition 1.** *Denote by  $P_n = E[p_{\underline{m}_n}(X)p_{\underline{m}_n}(X)']$ , and define*

$$\begin{aligned} \mu_{m_n} &\equiv \int E[|\delta(V, t)|^2 p_{\underline{m}_n}(X)' P_n^{-1} p_{\underline{m}_n}(X)] \varpi(t) dt \quad \text{and} \\ \varsigma_{m_n} &\equiv \left( \int \int \left\| P_n^{-1/2} E[\delta(V, s) \overline{\delta(V, t)} p_{\underline{m}_n}(X) p_{\underline{m}_n}(X)'] P_n^{-1/2} \right\|_F^2 \varpi(s) \varpi(t) ds dt \right)^{1/2}. \end{aligned}$$



Here, we use the notation  $\bar{\phi}$  for the complex conjugate of a function  $\phi$ , and  $\|\cdot\|_F$  to denote the Frobenius norm.

**Assumption 3.** *There exists some constant  $C > 0$  such that  $E[|\int \delta(V, t)\varpi(t)dt|^2|X] \geq C$ .*

Assumption 3 ensures that the conditional variance of  $\int \delta(V, t)\varpi(t)dt$  is uniformly bounded away from zero. Assumptions of this type are commonly required to obtain asymptotic normality of series estimators (see Assumption 4 of Newey [1997] or Theorem 4.2 of Belloni et al. [2015]). As we show in the appendix, Assumption 3 implies  $\varsigma_{m_n} \geq C\sqrt{m_n}$ .

### 2.3.2 Testing functional form restrictions

We now present conditions that are sufficient to provide the test's asymptotic distribution under the null hypothesis  $H_{\text{mod}}$ . To do so, let us introduce the norm  $\|\phi\|_{\varpi} = (\int E|\phi(X, t)|^2\varpi(dt))^{1/2}$  and the linear sieve space  $\Phi_n \equiv \{\phi : \phi(\cdot) = \sum_{j=1}^{m_n} \beta_j p_l(\cdot)\}$ . Moreover,  $\|\cdot\|$  and  $\|\cdot\|_{\infty}$ , respectively, denote the Euclidean norm and the supremum norm. Let us introduce  $A_n = \int E[(\mathcal{F}_g q_{k_n})(X, t)(\mathcal{F}_g q_{k_n})(X, t)']\varpi(t)dt$  and its empirical analog  $\hat{A}_n = n^{-1} \int \mathbf{F}_n(t)\mathbf{F}_n(t)'\varpi(t)dt$  (see also Example 4).

**Assumption 4.** *(i) For any p.d.f.  $f_B$  satisfying  $\varphi = \mathcal{F}_g f_B$  there exists  $\Pi_{k_n} f_B \in \mathcal{B}_n$  such that  $n\|\mathcal{F}_g(\Pi_{k_n} f_B - f_B)\|_{\varpi}^2 = o(\sqrt{m_n})$ . (ii) There exists  $\Pi_{m_n} \varphi \in \Phi_n$  such that  $n\|\Pi_{m_n} \varphi - \varphi\|_{\varpi}^2 = o(\sqrt{m_n})$  and  $\|\Pi_{m_n} \varphi - \varphi\|_{\infty} = O(1)$ . (iii) It holds  $k_n = o(\sqrt{m_n})$ . (iv) It holds  $A_n^- = O(1)$  and  $P(\text{rank}(A_n) = \text{rank}(\hat{A}_n)) = 1 + o(1)$ . (v) There exists a constant  $C > 0$  such that  $\sum_{l \geq 1} (\int_{\mathbb{R}^{d_b}} \phi(b)q_l(b)db)^2 \leq C \int_{\mathbb{R}^{d_b}} \phi^2(b)db$  for all square integrable functions  $\phi$ .*

Assumption 4 (i) is a requirement on the sieve approximation error for all functions  $f_B$  that belong to the identified set  $\mathcal{I}_g \equiv \{f : f \text{ is a p.d.f. with } \varphi = \mathcal{F}_g f\}$ . This condition ensures that the bias for estimating any  $f_B$  in the identified set  $\mathcal{I}_g$  is asymptotically negligible. Assumption 4 (ii) determines the sieve approximation error for the function  $\varphi$ . Consider the linear case and let  $\|\mathcal{F}(\Pi_{k_n} f_B - f_B)\|_{\varpi} = O(k_n^{-s/d_x})$  for some constant  $s > 0$ , then Assumptions 4 (i) and (iii) are satisfied if  $m_n \sim n^{\zeta}$  and  $k_n \sim n^{\kappa}$  where  $d_x(1 - \zeta/2)/(2s) < \kappa < \zeta/2$ .<sup>5</sup> We thus require  $\zeta > 2d_x/(2s + d_x)$ , so  $s$  has to increase with dimension  $d_x$ , which reflects a curse of dimensionality. In this case, Assumption 4 (ii) automatically holds if  $\|\Pi_{m_n} \varphi - \varphi\|_{\varpi} = O(m_n^{-s/d_x})$  and we may choose  $\kappa$  to balance variance and bias, i.e.,  $\kappa = d_x/(2s + d_x)$ .<sup>6</sup> For further discussion and examples of sieve bases, we refer to Chen [2007]. Assumption 4 (iv) ensures that the sequence of generalized inverse matrices is bounded and imposes a rank condition. This condition is

<sup>5</sup>We use the notation  $a_n \sim b_n$  for  $cb_n \leq a_n \leq Cb_n$  given two constant  $c, C > 0$  and all  $n \geq 1$ .

<sup>6</sup>This choice of  $k_n$  corresponds indeed to the optimal smoothing parameter choice in nonparametric random coefficient model if  $s = r + (d_x - 1)/2$  where  $r$  corresponds to the smoothness of  $f_B$  (see Hoderlein et al. [2010] in case of kernel density estimation).

sufficient and necessary for convergence in probability of generalized inverses of random matrices with fixed dimension (see Andrews [1987] for generalized Wald tests). Assumption 4 (v) is satisfied if  $\{q_l\}_{l \geq 1}$  forms a Riesz basis in  $L^2(\mathbb{R}^{d_b}) \equiv \{\phi : \int_{\mathbb{R}^{d_b}} \phi^2(s) ds < \infty\}$ . The following result establishes asymptotic normality of our standardized test statistic.

**Theorem 2.1.** *Let Assumptions 1–4 hold with  $\delta(V, t) = \exp(itY) - (\mathcal{F}_g f_B)(X, t)$ . Then, under  $H_{mod}$  we obtain*

$$(\sqrt{2}\varsigma_{m_n})^{-1}(n S_n - \mu_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1).$$

**Remark 2.1** (Estimation of Critical Values). *The asymptotic results of the previous theorem depends on unknown population quantities. As we see in the following, the critical values can be easily estimated. We define  $\delta_n(V, t) = \exp(itY) - (\mathcal{F}_g \hat{f}_{B_n})(X, t)$ , and*

$$\sigma_n(s, t) = \left( \delta_n(V_1, s) \overline{\delta_n(V_1, t)}, \dots, \delta_n(V_n, s) \overline{\delta_n(V_n, t)} \right)'$$

We replace  $\mu_{m_n}$  and  $\varsigma_{m_n}$ , respectively, by the estimators

$$\hat{\mu}_{m_n} = \int \text{tr} \left( (\mathbf{X}'_n \mathbf{X}_n)^{-1/2} \mathbf{X}'_n \text{diag}(\sigma_n(t, t)) \mathbf{X}_n (\mathbf{X}'_n \mathbf{X}_n)^{-1/2} \right) \varpi(t) dt$$

and

$$\hat{\varsigma}_{m_n} = \left( \int \int \left\| (\mathbf{X}'_n \mathbf{X}_n)^{-1/2} \mathbf{X}'_n \text{diag}(\sigma_n(s, t)) \mathbf{X}_n (\mathbf{X}'_n \mathbf{X}_n)^{-1/2} \right\|_F^2 \varpi(s) \varpi(t) ds dt \right)^{1/2}.$$

**Proposition 2.** *Under the conditions of Theorem 2.1, we obtain*

$$\varsigma_{m_n} \hat{\varsigma}_{m_n}^{-1} = 1 + o_p(1) \quad \text{and} \quad \hat{\mu}_{m_n} = \mu_{m_n} + o_p(\sqrt{m_n}).$$

The asymptotic distribution of our standardized test statistic remains unchanged if we replace  $\mu_{m_n}$  and  $\varsigma_{m_n}$  by estimators introduced in the last remark. This is summarized in following corollary, which follows immediately from Theorem 2.1 and Proposition 2.

**Corollary 2.1.** *Under the conditions of Theorem 2.1, we obtain*

$$(\sqrt{2}\hat{\varsigma}_{m_n})^{-1}(n S_n - \hat{\mu}_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1).$$

An alternative way to obtain critical values is the bootstrap which, for testing nonlinear functionals in nonparametric instrumental regression, was considered by Chen and Pouzo [2015]. In our situation, the critical values can be easily estimated and the finite sample properties of our testing procedure are promising, thus we do not elaborate bootstrap procedures here. In

the following example, we illustrate our sieve minimum distance approach for estimating  $f_B$  in the case of linearity of  $g$ .

**Example 7 (Linear Case).** *Let  $g$  be linear and recall that in this case the integral transform  $\mathcal{F}_g$  coincides with the Fourier transform  $\mathcal{F}$ . For the sieve space  $\mathcal{B}_n$ , we consider as basis functions Hermite functions given by*

$$q_l(x) = \frac{(-1)^l}{\sqrt{2^l l!} \sqrt{\pi}} \exp(x^2/2) \frac{d^l}{dx^l} \exp(-x^2).$$

*These functions form an orthonormal basis of  $L^2(\mathbb{R})$ . Hermite functions are also eigenfunctions of the Fourier transform with*

$$(\mathcal{F}q_l)(\cdot) = \sqrt{2\pi} (-i)^{-l} q_l(\cdot).$$

*Let us introduce the notation  $\tilde{q}_l(\cdot) \equiv (-i)^{-l} q_l(\cdot)$  and  $\mathbf{X}_n(t) = (\tilde{q}_{k_n}(tX_1)', \dots, \tilde{q}_{k_n}(tX_n)')$ . Thus, the estimator of  $f_B$  given in (2.7) simplifies to  $\hat{f}_{B_n}(\cdot) = q_{k_n}(\cdot)' \hat{\beta}_n$  where*

$$\hat{\beta}_n = \min_{\beta \in \mathbb{R}^{k_n}} \sum_{j=1}^n \int |\hat{\varphi}_n(X_j, t) - \tilde{q}_{k_n}(tX_j)' \beta|^2 \varpi(t) dt. \quad (2.10)$$

*An explicit solution of (2.10) is given by*

$$\hat{\beta}_n = \left( \int \mathbf{X}_n(t)' \mathbf{X}_n(t) \varpi(t) dt \right)^{-1} \int \mathbf{X}_n(t)' \Phi_n(t) \varpi(t) dt$$

*where  $\Phi_n(t) = (\hat{\varphi}_n(X_1, t), \dots, \hat{\varphi}_n(X_n, t))'$ . We emphasize that under the previous assumptions, the matrix  $\int \mathbf{X}_n(t)' \mathbf{X}_n(t) \varpi(t) dt$  will be nonsingular with probability approaching one.*

### 2.3.3 Testing degeneracy under the random coefficient specification for the model

For testing degeneracy, Theorem 2.1 is not directly applicable as the required sieve approximation error in Assumption 4 (i) is here not satisfied in general. In contrast, we will impose an approximation condition on the function  $\tilde{g}(x, t, b) \equiv \exp(itg_2(x, b))$  where  $b$  belongs to the parameter space  $\mathcal{B}_2$ .

Let us introduce a  $(k_n \cdot l_n)$ -dimensional vector valued function  $\chi_n$  given by  $\chi_n(x, t) = (\mathcal{F}_{g_1} q_{k_n})(x, t) \otimes \tilde{p}_{l_n}(x, t)$ , where  $\otimes$  denotes the Kronecker product and  $\tilde{p}_{l_n}$  is a  $l_n$ -dimensional vector of complex valued basis functions used to approximate  $\tilde{g}(\cdot, \cdot, b)$ . For instance, if  $g_2(x, b) = \phi(x)\psi(b)$  then approximation conditions can be easily verified due to  $\tilde{g}(x, t, b) = \sum_{l \geq 0} \tilde{p}_l(x, t) \psi(b)^l$  where  $\tilde{p}_l(x, t) = (it\phi(x))^l / l!$ . Let us introduce  $\mathcal{A}_n = \int E[\chi_n(X, t) \chi_n(X, t)'] \varpi(t) dt$  and its empirical analog  $\widehat{\mathcal{A}}_n = n^{-1} \int \sum_{j=1}^n \chi_n(X_j, t) \chi_n(X_j, t)' \varpi(t) dt$ . Recall that  $\mathcal{B}_{-2, n} = \{\phi(b) =$

$\sum_{l=1}^{k_n} \beta_l q_l(b)$  for  $b \in \mathbb{R}^{d_{b_2}}$  where  $d_{b_2}$  denotes the dimension of  $b_2$  and let  $\mathcal{G}_{2,n} = \{\phi(x, t) = \sum_{l=1}^{l_n} \beta_l \tilde{p}_l(x, t)\}$ .

**Assumption 5.** (i) The hypothesis  $H_{odd}$  holds. (ii) The set of parameters  $b_2$  satisfying (2.5) belongs to a compact parameter space  $\mathcal{B}_2 \subset \mathbb{R}^{d_{b_2}}$ . (iii) For any  $b \in \mathcal{B}_2$  there exists  $\Pi_{l_n} \tilde{g}(\cdot, \cdot, b) \in \mathcal{G}_{2,n}$  satisfying  $n \|\Pi_{l_n} \tilde{g}(\cdot, \cdot, b) - \tilde{g}(\cdot, \cdot, b)\|_{\varpi}^2 = o(\sqrt{m_n})$ . (iv) For any p.d.f.  $f_{B_{-2}}$  satisfying (2.5) there exists  $\Pi_{k_n} f_{B_{-2}} \in \mathcal{B}_{-2,n}$  such that  $n \|\mathcal{F}_{g_1}(\Pi_{k_n} f_{B_{-2}} - f_{B_{-2}})\|_{\varpi}^2 = o(\sqrt{m_n})$ . (v) It holds  $k_n l_n = o(\sqrt{m_n})$ . (vi) It holds  $\mathcal{A}_n^- = O(1)$  and  $P(\text{rank}(\mathcal{A}_n) = \text{rank}(\widehat{\mathcal{A}}_n)) = 1 + o(1)$ . (vii) There exists a constant  $C > 0$  such that  $\sum_{l, l' \geq 1} \langle \mathcal{F}_{g_1} q_l \cdot \tilde{p}_{l'}, \phi \rangle^2 \leq C \|\phi\|_{\varpi}^2$  for all functions  $\phi$  with  $\|\phi\|_{\varpi} < \infty$ .

Assumption 5 (i) states the maintained hypothesis of an additive structure of  $g$  given in equation (2.1). Assumption 5 (iii) states an asymptotic condition of the sieve approximation error for  $\tilde{g}(\cdot, \cdot, b)$  for any  $b$  in the parameter space  $\mathcal{B}_2$ . By doing so, we impose regularity conditions on the integral transform  $\mathcal{F}_{g_2}$  of the Dirac measure at  $b$  but not on the Dirac measure itself. For instance, if again  $g_2(x, b) = \phi(x)\psi(b)$  and  $\tilde{p}_l(x, t) = (it\phi(x))^l/l!$  for  $l \geq 1$  then  $\|\Pi_{l_n} \tilde{g}(\cdot, \cdot, b) - \tilde{g}(\cdot, \cdot, b)\|_{\varpi} \leq C/(l_n + 1)!$  for some constant  $C > 0$ , provided that  $E[\phi^{l_n}(X)] \psi^{l_n}(b) \int t^{l_n} \varpi(t) dt$  is bounded. Assumption 5 (v) restricts the magnitude of  $k_n$  also relative to the dimension parameter  $l_n$ , which is not too restrictive as the dimension  $k_n$  is used to approximate a lower dimensional p.d.f. than in Theorem 2.1. Assumption 5 (iii) requires an appropriate sieve approximation error only for any nondegenerate p.d.f.  $f_{B_{-2}}$  satisfying (2.5). Assumption 5 (vi) and (vii), respectively, are closely related to Assumption 4 (iv) and (v).

**Theorem 2.2.** *Let Assumptions 1–3, 4 (ii), and 5 be satisfied with  $\delta(V, t) = \exp(itY) - (\mathcal{F}_{g_1} f_{B_{-2}})(X, t) \tilde{g}(X, t, b_2)$ . Then, under  $H_{deg}$  we obtain*

$$(\sqrt{2\varsigma_{m_n}})^{-1}(n S_n - \mu_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1).$$

The critical values can be estimated as in Remark 2.1 but where now  $\delta_n(V, t) = \exp(itY) - (\mathcal{F}_{g_1} \widehat{f}_{B_{-2n}})(X, t) \tilde{g}(X, t, \widehat{b}_{2n})$ . The following result shows that, by doing so, the asymptotic distribution of our standardized test statistic remains unchanged. This corollary follows directly from Theorem 2.3 and the proof of Proposition 2; hence we omit its proof.

**Corollary 2.2.** *Under the conditions of Theorem 2.3 it holds*

$$(\sqrt{2\widehat{\varsigma}_{m_n}})^{-1}(n S_n - \widehat{\mu}_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1).$$

**Remark 2.2** (Comparison to Andrews [2001]). *It is instructive to compare our setup and*

results to Andrews [2001], who considers the random coefficient model:

$$Y = B_0 + B_1 X_1 + (b_2 + \tau \tilde{B}_2) X_2,$$

where  $E[B_0 \cdot B_1 | X] = 0$ ,  $B_1$  is independent of  $\tilde{B}_2$ , and  $E[B_1 | X] = E[\tilde{B}_2 | X] = 0$ . In this model, degeneracy of the second random coefficient is equivalent to  $\tau = 0$  and degeneracy fails if  $\tau > 0$ . So under  $H_{deg}$  the parameter  $\tau$  is on the boundary of the maintained hypothesis with  $\tau \in [0, \infty)$ .

In contrast, we rely in this paper on independence of  $B$  to  $X$  under the maintained hypothesis. In this case, the hypothesis of degeneracy is equivalent to a conditional characteristic function equation as explained in Example 2 and which is not possible given the assumptions of Andrews [2001]. This is why in our framework we automatically avoid the boundary problem that is apparent in Andrews [2001].

### 2.3.4 Testing degeneracy under additive separability alone

We now establish the asymptotic distribution of our test of degeneracy based on separability but not full knowledge of  $g_1$  (see Examples 3 and 6). We introduce the function  $h(\cdot, t) = E[\exp(it(Y - g_2(X, b_2))) | X_1 = \cdot]$  and a linear sieve space  $\mathcal{H}_n \equiv \{\phi : \phi(x_1) = \sum_{j=1}^{k_n} \beta_j p_l(x_1) \text{ for } x_1 \in \mathbb{R}^{d_{x_1}}\}$  where  $d_{x_1}$  denotes the dimension of  $X_1$ . The series least squares estimator of  $h$  is denoted by  $\hat{h}_n(\cdot) = p_{k_n}(\cdot)' (\mathbf{X}'_{1n} \mathbf{X}_{1n})^{-1} \mathbf{X}'_{1n} \mathbf{U}_n$  where  $\mathbf{U}_n = (\exp(it(Y_1 - g_2(X_1, \hat{b}_{2n}))), \dots, \exp(it(Y_n - g_2(X_n, \hat{b}_{2n}))))'$  and  $\hat{b}_{2n}$  denotes an estimator of  $b_2$ . Recall the notation  $\tilde{g}(x, t, b) \equiv \exp(itg_2(x, b))$  for  $b \in \mathcal{B}_2$ . Below we denote the vector of partial derivatives of  $\tilde{g}$  with respect to  $b$  by  $\tilde{g}_b$ .

**Assumption 6.** (i) The hypothesis  $H_{add}$  holds, where  $g_1$  need not to be known except that it does not depend on  $X_2$ . (ii) There exists  $\Pi_{k_n} h \in \mathcal{H}_n$  such that  $n \|\Pi_{k_n} h - h\|_{\varpi}^2 = o(\sqrt{m_n})$ . (iii) The parameter  $b_2$  is point identified and belongs to the interior of a compact parameter space  $\mathcal{B}_2 \subset \mathbb{R}^{d_{b_2}}$ . (iv) There exists an estimator  $\hat{b}_{2n}$  such that  $\sqrt{n}(\hat{b}_{2n} - b_2) = O_p(1)$  (v) The function  $\tilde{g}$  is partially differentiable with respect to  $b$  and  $\int E \sup_{b \in \mathcal{B}_2} \|\tilde{g}_b(X, t, b)\|_{\varpi}^2 dt < \infty$ . (vi) It holds  $k_n = o(\sqrt{m_n})$ .

Assumption 6 (ii) determines the required asymptotic behavior of the sieve approximation bias for estimating  $h$ . This condition ensures that the bias for estimating the function  $h$  is asymptotically negligible but does not require undersmoothing of the estimator  $\hat{h}_n$ . To see this, let  $\|\Pi_{k_n} h - h\|_{\varpi} = O(k_n^{-s/d_{x_1}})$  for some constant  $s > 0$ . Assumptions 6 (ii) and (vi) are satisfied if  $m_n \sim n^{\zeta}$  and  $k_n \sim n^{\kappa}$  where  $d_{x_1}(1 - \zeta/2)/(2s) < \kappa < \zeta/2$ . We thus require  $\zeta > 2d_{x_1}/(2s + d_{x_1})$  and we may choose  $\kappa$  to balance variance and bias, i.e.,  $\kappa = d_{x_1}/(2s + d_{x_1})$ . In this case, Assumption 4 (ii) automatically holds if  $\|\Pi_{m_n} \varphi - \varphi\|_{\varpi} = O(m_n^{-s/d_x})$  and  $2d_{x_1} \geq d_x$ . Under a partially linear structure  $H_{part-lin}$ , Assumption 6 (iv) is automatically satisfied if  $\hat{b}_{2n}$

coincides with the nonlinear least squares estimator. If  $g_2$  is linear, Assumption 6 (iv) holds true if  $E\|X\|^2 < \infty$  and  $\int t^2 \varpi(t)dt < \infty$ .

**Theorem 2.3.** *Let Assumptions 1–3, 4 (ii), and 6 hold, with  $\delta(V, t) = \exp(itY) - h(X_1, t)\tilde{g}(X, t, b_2)$ . Then, under  $H_{deg}$  we obtain*

$$(\sqrt{2}\zeta_{m_n})^{-1}(n S_n - \mu_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1).$$

The critical values can be estimated as in Remark 2.1 but where now  $\delta_n(V, t) = \exp(itY) - \widehat{h}_n(X_1, t) \exp(itg_2(X, \widehat{b}_{2n}))$ . The following result shows that, by doing so, the asymptotic distribution of our standardized test statistic remains unchanged. This corollary follows directly from Theorem 2.3 and the proof of Proposition 2; hence we omit its proof.

**Corollary 2.3.** *Under the conditions of Theorem 2.3 it holds*

$$(\sqrt{2}\widehat{\zeta}_{m_n})^{-1}(n S_n - \widehat{\mu}_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1).$$

## 2.4 Consistency against a fixed alternative

In the following, we establish consistency of our test when the difference of restricted and unrestricted conditional characteristic functions does not vanish for all random parameters  $B$ . In case of testing functional form restrictions, this is equivalent to a failure of the null hypothesis  $H_{mod}$ , i.e.,  $P(Y \neq g(X, B) \text{ for all random parameters } B) > 0$ . A deviation of conditional characteristic functions can be also caused by alternative models with a different structural function (see Example 1). We only discuss the global power for testing functional form restrictions here, but the results for testing degeneracy follow analogously (of course, in this case we have to be more restrictive about the shape of  $g_1$  and  $g_2$  as discussed in Example 2). The next proposition shows that our test of functional form restrictions has the ability to reject a failure of the null hypothesis  $H_{mod}$  with probability one as the sample size grows to infinity.

**Proposition 3.** *Suppose that  $H_{mod}$  is false and let Assumptions 1–4 be satisfied. Consider a sequence  $(\gamma_n)_{n \geq 1}$  satisfying  $\gamma_n = o(n\zeta_{m_n}^{-1})$ . Then, we have*

$$P\left((\sqrt{2}\widehat{\zeta}_{m_n})^{-1}(n S_n - \widehat{\mu}_{m_n}) > \gamma_n\right) = 1 + o(1).$$

## 2.5 Asymptotic distribution under local alternatives

We now study the power of our testing procedure against a sequence of linear local alternatives that tends to zero as the sample size tends to infinity. First, we consider deviations from

the hypothesis of known functional form restriction. Under  $H_{\text{mod}}$ , the identified set in the nonseparable model (1.1) is given by  $\mathcal{I}_g = \{f : f \text{ is a p.d.f. with } \varphi = \mathcal{F}_g f\}$ . We consider the following sequence of local alternatives

$$\varphi = \mathcal{F}_g(f_B^* + \Delta\sqrt{\varsigma_{m_n}/n}), \quad (2.11)$$

for some function  $\Delta \in L^1(\mathbb{R}^{d_b}) \cap L^2(\mathbb{R}^{d_b})$  and  $f_B^*$  is a p.d.f. satisfying  $\|\mathcal{F}_g f_B^* - \varphi\|_{\varpi} \leq \|\mathcal{F}_g f - \varphi\|_{\varpi}$  for any p.d.f.  $f$ . Here, we assume that  $\Delta$  is such that  $f_B + \Delta\sqrt{\varsigma_{m_n}/n}$  does not belong to the identified set  $\mathcal{I}_g$ . The next result establishes asymptotic normality under (2.11) of the standardized test statistic  $S_n$  for testing functional form restrictions.

**Proposition 4.** *Let the assumptions of Theorem 2.1 be satisfied. Then, under (2.11) we obtain*

$$(\sqrt{2}\widehat{\varsigma}_{m_n})^{-1}(n S_n - \widehat{\mu}_{m_n}) \xrightarrow{d} \mathcal{N}\left(2^{-1/2}\|\mathcal{F}_g \Delta\|_{\varpi}^2, 1\right).$$

As we see from Proposition 4, our test can detect linear alternatives at the rate  $\sqrt{\varsigma_{m_n}/n}$ . Results for testing degeneracy follow similarly. In the following, we thus study deviations from the hypothesis of degeneracy only under the maintained hypothesis  $H_{\text{lin}} : Y = B_0 + B'_1 X_1 + B'_2 X_2$ . Under the maintained hypothesis of linearity, any deviation between the conditional characteristic functions is equivalent to a failure of a degeneracy of the random coefficients  $B_2$ . Let us denote  $B_{deg} \equiv (B_1, b_2)$  with associated p.d.f.  $f_{B_{deg}}$ . We consider the following sequence of linear local alternatives

$$f_B = f_{B_{deg}} + \Delta\sqrt{\varsigma_{m_n}/n}, \quad (2.12)$$

for some density function  $\Delta \in L^1(\mathbb{R}^{d_b}) \cap L^2(\mathbb{R}^{d_b})$  which is not degenerate at  $b_2$ . Applying the Fourier transform to equation (2.12) yields

$$E[\exp(itX'B)|X] = E[\exp(it(B_0 + X'_1 B_1))|X] \exp(itX'_2 b_2) + \int \exp(itX's)\Delta(s)ds\sqrt{\varsigma_{m_n}/n}.$$

The next result establishes asymptotic normality under (2.12) for the standardized test statistic  $S_n$  for testing degeneracy. This corollary follows by similar arguments used to establish Proposition 4 and hence we omit the proof.

**Corollary 2.4.** *Let the assumptions of Theorem 2.3 be satisfied. Then, under (2.12) we obtain*

$$(\sqrt{2}\widehat{\varsigma}_{m_n})^{-1}(n S_n - \widehat{\mu}_{m_n}) \xrightarrow{d} \mathcal{N}\left(2^{-1/2}\|\mathcal{F}\Delta\|_{\varpi}^2, 1\right).$$

### 3 Extensions

In this section, we show that our testing procedures can be extended to two different models. First, we consider the class of heterogeneous binary response models. Second, we discuss an extension of linear random coefficient models to system of equations. In both cases, we again discuss testing functional form restrictions and testing degeneracy of some random coefficients separately.

#### 3.1 Binary Response Models

We consider the binary response model

$$Y = 1\{g(X, B) < Z\}, \quad (3.1)$$

where, besides the dependent variable  $Y$  and covariates  $X$ , a *special regressor*  $Z$  is observed as well. In the following, we assume that  $(X, Z)$  is independent of  $B$ . In contrast to the previous section, the test in the binary response model is based on the difference of a partial derivative of the conditional success probability  $P(Y = 1|X, Z)$  and a restricted transformation of the p.d.f.  $f_B$ .

**Testing functional form restrictions.** In the binary response model (3.1), observe that

$$\begin{aligned} P[Y = 1|X = x, Z = z] &= \int 1\{z > g(x, b)\} f_B(b) db \\ &= \int_{-\infty}^z \int_{P_{x,s}} f_B(b) d\nu(b) ds, \end{aligned}$$

where  $\nu$  is the Lebesgue measure on the lower dimensional hyperplane  $P_{x,s} = \{b : g(x, b) = s\}$ . Consequently, it holds

$$\psi(x, z) \equiv \partial_z P[Y = 1|X = x, Z = z] = \int_{P_{x,z}} f_B(b) d\nu(b).$$

Again by considering conditional characteristic functions, the null hypothesis  $H_{\text{mod}}$  is equivalent to  $E[\exp(itY)|X] = E[\exp(itg(X, B))|X]$  for some random coefficient  $B$ . By using the above integral representation of  $\psi$  we equivalently obtain  $(\mathcal{F}\psi(X, \cdot))(t) = (\mathcal{F}_g f_B)(t, X)$  (recall the definition of the integral transform  $(\mathcal{F}_g f)(X, t) \equiv \int \exp(itg(X, b))f(b)db$ ). Due to technical reason, we invert the Fourier transform and conclude that equation (2.3) holds true with

$$\varepsilon(X, z) = \psi(X, z) - (\mathcal{F}^{-1}[(\mathcal{F}_g f_B)(X, \cdot)])(z).$$



In the case of a linear  $g$ , the random coefficient density  $f_B$  is thus identified through the Radon transform, see also Gautier and Hoderlein [2015].

To estimate the function  $\varepsilon$ , we replace  $\psi$  by a series least squares estimator. Let us introduce the matrix  $\mathbf{W}_n = (p_{\underline{m}_n}(X_1, Z_1), \dots, p_{\underline{m}_n}(X_n, Z_n))'$  where the basis function  $p_l$ ,  $l \geq 1$ , are assumed to be differentiable with respect to the  $(d_x + 1)$ -th entry. We estimate  $\psi$  by

$$\widehat{\psi}_n(x, z) = \partial_z p_{\underline{m}_n}(x, z)' (\mathbf{W}_n' \mathbf{W}_n)^{-1} \mathbf{Y}_n,$$

where  $\mathbf{Y}_n = (Y_1, \dots, Y_n)'$ . Consequently, we replace the function  $\varepsilon$  by

$$\widehat{\varepsilon}_n(X_j, z) = \widehat{\psi}_n(X_j, z) - (\mathcal{F}^{-1}[(\mathcal{F}_g \widehat{f}_{Bn})(X_j, \cdot)])(z),$$

where  $\widehat{f}_{Bn}$  is the sieve minimum distance estimator given by

$$\widehat{f}_{Bn} \in \arg \min_{f \in \mathcal{B}_n} \left\{ \sum_{j=1}^n \int |\widehat{\psi}_n(X_j, z) - (\mathcal{F}^{-1}[(\mathcal{F}_g f)(X_j, \cdot)])(z)|^2 \varpi(z) dz \right\} \quad (3.2)$$

and  $\mathcal{B}_n = \{\phi(\cdot) = \sum_{l=1}^{k_n} \beta_l q_l(\cdot)\}$ . Our test statistic is  $S_n = n^{-1} \sum_{j=1}^n \int |\widehat{\varepsilon}_n(X_j, z)|^2 \varpi(z) dz$  where, in this section,  $\varpi$  is an integrable weighting function on the support of  $Z$ .

We introduce an  $m_n$  dimensional linear sieve space  $\Psi_n \equiv \{\phi : \phi(x, z) = \sum_{j=1}^{m_n} \beta_j p_j(x, z)\}$ . Let  $p_{\underline{m}_n}(X, Z)$  be a tensor-product of vectors of basis functions  $p_{\underline{m}_{n_1}}(X)$  and  $p_{\underline{m}_{n_2}}(Z)$  for integers  $m_{n_1}$  and  $m_{n_2}$  with  $m_n = m_{n_1} \cdot m_{n_2}$ . We assume that  $\partial_z p_{\underline{m}_{n_2}}(z) = (p_0(z), 2p_1(z), \dots, m_{n_2} p_{m_{n_2}-1}(z))'$ . Further, let  $\tau_l$  denote the squared integer that is associated with  $\partial_z p_l$ . In Definition 1,  $p_l(X)$  has to be replaced by  $\tau_l p_l(X, Z)$ . Let  $B_n = \int E[(\mathcal{F}^{-1}[(\mathcal{F}_g q_{k_n})(X, \cdot)])(z) (\mathcal{F}^{-1}[(\mathcal{F}_g q_{k_n})(X, \cdot)])(z)'] \varpi(z) dz$ , which is denoted by  $\widehat{B}_n$  when the expectation is replaced by the sample mean.

**Assumption 7.** (i) The random vector  $(X, Z)$  is independent of  $B$ . (ii) For any p.d.f.  $f_B$  satisfying  $\mathcal{F}\psi = \mathcal{F}_g f_B$  there exists  $\Pi_{k_n} f_B \in \Psi_n$  such that  $n \|\psi - \mathcal{F}^{-1} \mathcal{F}_g \Pi_{k_n} f_B\|_{\varpi}^2 = o(\sqrt{m_n})$ . (iii) There exists  $\Pi_{m_n} \psi \in \Psi_n$  such that  $n \|\Pi_{m_n} \psi - \psi\|_{\varpi}^2 = o(\sqrt{m_n})$ . (iv) It holds  $B_n^- = O(1)$  and  $P(\text{rank}(B_n) = \text{rank}(\widehat{B}_n)) = 1 + o(1)$ . (v) It holds  $k_n = o(\sqrt{m_n})$  and  $m_n^2 (\log n) \sum_{l=1}^{m_n} \tau_l = o(n \lambda_n)$ .

Assumption 7 is similar to Assumption 4. Note that due to the partial derivatives of the basis functions we need to be more restrictive about the dimension parameter  $m_n$ , which is captured in Assumption 7 (iv). The following result establishes the asymptotic distribution of our test statistic under  $H_{\text{mod}}$  in the binary response model (3.1).

**Proposition 5.** Let Assumptions 2, 3, and 7 hold with  $\delta(Y, X, Z) = Y - \int 1\{Z > X'b\} f_B(b) db$ . Then, under  $H_{\text{mod}}$  we have

$$(\sqrt{2\varsigma_{m_n}})^{-1} (n S_n - \mu_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1).$$

The critical values can be estimated as in Remark 2.1 but where now  $\delta_n(Y, X, Z) = Y - \int 1\{Z \geq X'b\} \widehat{f}_{B_n}(b) db$  with the estimator  $\widehat{f}_{B_n}$  given in (3.2).

**Testing degeneracy.** To keep the presentation simple, we only consider the linear case in the following. Under  $H_{lin}$ , the binary response model (3.1) simplifies to

$$Y = 1\{X'B < Z\}. \quad (3.3)$$

$$\int \exp(itz) \psi(X, z) dz = \int \exp(itz) \psi(X_1, z - X_2'b_2) dz.$$

By nonsingularity of the Fourier transform, we conclude that  $H_{deg}$  is equivalent to equation (2.3) where

$$\varepsilon(X, z) = \psi(X, z) - \psi(X_1, z - X_2'b_2).$$

If  $\psi$  only depends on  $X_1$ , we consider the estimator  $\widehat{\psi}_{1n}(x_1, z) = \partial_z p_{k_n}(x_1, z)' (\mathbf{W}'_{n1} \mathbf{W}_{n1})^{-1} \mathbf{Y}_n$ , where  $\mathbf{W}_{n1} = (p_{k_n}(X_{11}, Z_1), \dots, p_{k_n}(X_{1n}, Z_n))'$ . We propose a minimum distance estimator of  $b_2$  given by

$$\widehat{b}_{2n} = \operatorname{argmin}_{\beta \in \mathcal{B}} \sum_{j=1}^n \int \left| \widehat{\psi}_n(X_j, t) - \widehat{\psi}_{1n}(X_{1j}, t - \beta' X_2) \right|^2 \varpi(t) dt. \quad (3.4)$$

Consequently, we estimate the function  $\varepsilon$  by  $\widehat{\varepsilon}_n(X_j, z) = \widehat{\psi}_n(X_j, z) - \widehat{\psi}_{1n}(X_{1j}, z - \widehat{b}_{2n}' X_{2j})$ .

**Proposition 6.** *Let Assumptions 2, 3, 7 (i), (iii), (v) with  $\delta(Y, X, Z) = Y - P(Y = 1|X_1, Z - X_2'b_2)$ , and  $H_{lin}$  hold true. Assume that  $n \int E|(\Pi_{k_n} \psi)(X_1, z) - \psi(X_1, z)|^2 \varpi(z) dz = o(\sqrt{m_n})$ . Then, under  $H_{deg}$  we have*

$$(\sqrt{2} \varsigma_{m_n})^{-1} (n S_n - \mu_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1).$$

The critical values can be estimated as in Remark 2.1 by replacing  $P(Y = 1|X_1, Z - X_2'b_2)$  by a series least squares estimator.

### 3.2 Application to Systems of Equations

In this subsection, we apply our testing procedure to systems of equations, i.e., situations in which the endogenous variable is not a scalar, but a vector. For simplicity, we consider in the

following only the bivariate case. Formally, we consider the model

$$Y = g(X, B), \tag{3.5}$$

for some function  $g$  and  $Y \in \mathbb{R}^2$ . Again the vector of random coefficients  $B = (B_0, B_1, B_2, B_3)$  is assumed to be independent of the covariates  $X$ .

**Testing functional form restrictions.** Null hypothesis  $H_{\text{mod}}$  is equivalent to equation (2.3) with

$$\varepsilon(X, t) = E[\exp(it'Y) - \exp(it'g(X, B))|X]$$

for some  $t \in \mathbb{R}^2$ . Our test of  $H_{\text{mod}}$  is now based on  $S_n \equiv n^{-1} \sum_{j=1}^n \int |\widehat{\varepsilon}_n(X_j, t)|^2 \varpi(t) dt$  where  $\widehat{\varepsilon}_n$  is the estimator of  $\varepsilon$  introduced in Example 4 but with a multivariate index  $t$  and  $\varpi$  being a weighting function on  $\mathbb{R}^2$ . Under a slight modification of assumptions required for Theorem 2.1, asymptotic normality of the standardized test statistic  $S_n$  follows under  $H_{\text{mod}}$ .

**Testing degeneracy.** In the partially linear case (i.e.,  $H_{\text{part-lin}}$  holds), the random coefficient model (3.5) simplifies to

$$\begin{aligned} Y_1 &= B_0 + B'_{11}X_1 + B'_{12}X_2 \\ Y_2 &= B_2 + B'_{31}X_1 + g_2(X_2, B_{32}). \end{aligned}$$

This model is identified if  $B_{32}$  is degenerate (see Hoderlein et al. [2014]). A test for degeneracy of  $H_{\text{deg}} : B_{32} = b$ , for some non-stochastic vector  $b$ , uses only the second equation, i.e.,

$$E[\exp(itY_2)|X] = E[\exp(it(B_2 + B'_{31}X_1))|X_1] \exp(itg_2(X_{32}, b)).$$

We can consequently use the testing methodology developed in Section 2.3.4.

## 4 Monte Carlo Experiments

In this section, we study the finite-sample performance of our test by presenting the results of a Monte Carlo simulation study. The experiments use a sample size of 500 and there are 1000 Monte Carlo replications in each experiment. As throughout the paper, we structure this section again in a part related to testing functional form restrictions, and a part related to testing degeneracy.

## 4.1 Testing Functional Form Restrictions

In each experiment, we generate realizations of regressors  $X$  from  $X \sim \mathcal{N}(0, 2)$  and random coefficients  $B = (B_1, B_2)'$  from  $B \sim \mathcal{N}(0, A)$  where

$$A = \begin{pmatrix} 1 & 1/2 \\ 1/2 & 1 \end{pmatrix}.$$

We simulate a random intercept  $B_0 \perp (B_1, B_2)$  according to the standard normal distribution. Realizations of the dependent variable  $Y$  are generated either by the linear model

$$Y = \eta B_0 + X B_1, \tag{4.1}$$

the quadratic model

$$Y = c_1(\eta B_0 + X B_1 + X^2 B_2), \tag{4.2}$$

or the nonlinear model

$$Y = c_2(\eta B_0 + X B_1 + \sqrt{|X|} B_2), \tag{4.3}$$

where the constant  $\eta$  is either  $\sqrt{0.5}$  or 1. Here, the normalization constants  $c_1$  and  $c_2$  ensure that the dependent variables in models (4.1)–(4.3) have the same variance.<sup>7</sup> Note that the random coefficient density  $f_B$  is neither point identified in model (4.2) nor in model (4.3). However, recall that even if the model is not point identified under the maintained hypothesis, our testing procedure may still be able to detect certain failures of the null hypothesis, in particular if they arise from differences in conditional moments. Consider, for example, testing linearity in the heterogeneous QUAIDS model (4.2), where the first two conditional moments yield  $E[B_2] = 0$  and  $Var(B_2) = 0$ . Consequently,  $P(\int |\varepsilon(X, t)|^2 \varpi(t) dt \neq 0) > 0$  if and only if  $P(B_2 \neq 0) > 0$ . We also observe in the finite sample experiment that our testing procedure is able to detect such deviations.

The test is implemented using Hermite functions, and uses the standardization described in Remark 2.1. When (4.1) is the true model, we estimate the random coefficient density as described in Example 7, where we make use of the fact that the Hermite functions are eigenfunctions of the Fourier transform. If (4.2) is the true model, the integral transform  $\mathcal{F}_g$  is computed using numerical integration. In both cases, the weighting function  $\varpi$  coincides with the standard normal p.d.f.. If (4.1) is the correct model, we use  $k_n = 4 (= 3 + 2)$  Hermite functions to estimate the density of the bivariate random coefficients  $(B_0, B_1)$  and let  $m_n = 9$ .

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<sup>7</sup>This normalization ensures that large empirical rejection probabilities are not only driven by a large variance of the alternative models (see, for instance, Blundell and Horowitz [2007]).

Null Model $H_{\text{mod}}$	Alt. Model True DGP	$\eta$	Empirical Rejection probabilities at level		
			0.010	0.050	0.100
(4.1)		0.5	0.003	0.027	0.076
(4.2)			0.008	0.034	0.072
(4.1)	(4.2)	1	0.698	0.911	0.958
(4.1)	(4.3)		0.178	0.491	0.683
(4.2)	(4.1)		0.714	0.928	0.980
(4.2)	(4.3)		0.864	0.982	0.994
(4.1)		1	0.007	0.047	0.127
(4.2)			0.007	0.058	0.154
(4.1)	(4.2)		0.491	0.804	0.907
(4.1)	(4.3)		0.102	0.368	0.615
(4.2)	(4.1)		0.557	0.875	0.970
(4.2)	(4.3)		0.558	0.905	0.978

Table 1: Rows 1,2,7,8 depict the empirical rejection probabilities if  $H_{\text{mod}}$  holds true, the rows 3–6 and 9–12 show the finite sample power of our tests against various alternatives. The first column states the null model while the second shows the alternative model and is left empty if the null model is the correct model. Column 3 specifies the noise level of the data generating process. Columns 4–6 depict the empirical rejection probabilities for different nominal levels.

If (4.2) is the correct model, we have an additional dimension which accounts for the nonlinear part. Here, the choice of Hermite basis functions is  $k_n = 7 (= 3 + 2 \cdot 2)$  with  $m_n = 12$  if  $\eta = \sqrt{0.5}$  and  $k_n = 9 (= 3 + 2 \cdot 3)$  with  $m_n = 16$  if  $\eta = 1$ . We thus increase the dimension parameters  $k_n$  and  $m_n$  as the noise level  $\eta$  becomes larger, i.e., the model becomes more complex. Note that  $m_n$  could be any integer larger than  $\text{const.} \times k_n^2$  that is smaller than  $n^{1/2}$  (up to logs). In practice, we let  $k_n$  such that it minimizes the value of the test statistic. I.e., if  $s(k_n, m_n)$  denotes the value of the test statistic, a guideline for parameter choice in practice is given by the minimum-maximum principle  $\min_{1 \leq k_n < n^{1/4}} \max_{k_n^2 < m_n < n^{1/2}} \{s(k_n, m_n)\}$ .

The empirical rejection probabilities of our tests are shown in Table 1 at nominal levels 0.010, 0.050, and 0.100. We also note that the models are normalized and hence, the null and alternative have the same variance. The differences between the nominal and empirical rejection probabilities are small under the correct functional form restrictions, as is obvious from rows 1, 2, 7, and 8. Comparing the empirical rejections probabilities in rows 3–6 and 9–12, we see that our tests become less powerful as the parameter  $\eta$  increases, as was to be expected. On the other hand, we observe from this table that our tests have power to detect nonlinear alternatives even in cases where the model under the maintained hypothesis is not identified. This is in line with our observation that these alternatives imply deviations between the restricted and unrestricted characteristic functions. Comparing rows 3, 9 with 4, 10 in Table 1, we observe

that our test rejects the quadratic model (4.2) more often than the nonlinear model (4.3). From rows 5, 11 and 6, 12 we see that our test rejects the nonlinear model (4.3) slightly more often than the linear model (4.1).

We have also tried different data generating processes, such as a cubic polynomial with random coefficients. In this case, our test of linearity led to empirical rejection probabilities which were close to one for all nominal levels considered and hence these results are not reported here. Regarding consistency, we also conducted experiments with larger sample sizes. In particular, we saw that the slight tendency of our test statistic to under-reject for small  $\eta$ , see in Table 1 in rows 1 and 2, diminishes as we increase the sample size to  $n = 1000$ . Not surprisingly, when  $n = 1000$  also the empirical rejection probabilities in rows 3–6 and 9–12 increase.

## 4.2 Testing Degeneracy

In each experiment, we generate realizations of  $X$  from  $X \sim \mathcal{N}(0, A)$  and random coefficients  $B = (B_1, B_2)'$  from  $B \sim \mathcal{N}(0, A_\rho)$ , where

$$A = \begin{pmatrix} 1 & .5 \\ .5 & 1 \end{pmatrix} \quad \text{and} \quad A_\rho = \begin{pmatrix} 2 & \rho \\ \rho & 2 \end{pmatrix},$$

for some constant  $\rho > 0$ , which varies in the experiments. Further, we generate the dependent variable  $Y$  as

$$Y = B_0 + B_1 X_1 + X_2,$$

if the null hypothesis  $H_{\text{deg}}$  holds. For the alternative, we generate the dependent variable  $Y$  using

$$Y = B_0 + B_1 X_1 + \eta B_2 X_2,$$

for some constant  $\eta > 0$ , which varies in the simulations below.

The test is implemented as described in Example 6 with B-splines, and uses the standardization described in Remark 2.1 with  $\delta_n(V, t) = \exp(itY) - \widehat{h}_n(X_1, t) \exp(itg_2(X, \widehat{b}_{2n}))$ . This means that we use the more general test that allows for a nearly arbitrary specification in the remaining model  $Y - g_2(X_2, b_2)$ . We focus in the simulation on this specification, because it has arguably less power than the more specific one that imposes in addition the linear random coefficients structure. However, as will be evident from the results below, this test already has very good power properties, implying that separating the term involving the fixed coefficient turns out to already be a powerful device in testing. To estimate the restricted conditional characteristic function, we use B-splines of order 2 with one knot (hence,  $k_n = 4$ ), and for the

<i>Alt. Model</i>		<i>Empirical Rejection probabilities at level</i>		
$\eta$	$\rho$	<i>0.010</i>	<i>0.050</i>	<i>0.100</i>
		0.003	0.049	0.108
0.75	1	0.033	0.177	0.352
1		0.142	0.479	0.837
1.25		0.378	0.738	0.876
0.75	1.5	0.118	0.396	0.643
1		0.367	0.741	0.837
1.25		0.627	0.904	0.967

Table 2: *The first row depicts the empirical rejection probabilities under degeneracy of the coefficient of  $X_2$ , the rows 2–7 show the finite sample power of our tests against various alternatives. Column 1 depicts the value of  $\eta$  in the correct model and is empty if the null model is correct. Column 2 specifies the covariance of  $B_1$  and  $B_2$  for the alternative models. Columns 4–6 depict the empirical rejection probabilities for different nominal levels.*

unrestricted one a tensor-product of this B-spline basis functions and a quadratic polynomial (hence,  $m_n = 12$ ). In practice, we may employ the minimum-maximum principle for parameter choice, as described in the previous subsection.

The empirical rejection probabilities for testing degeneracy are shown in Table 2 at nominal levels 0.010, 0.050, and 0.100. Again we normalize the models to ensure that the null and alternative have the same variance. The differences between the nominal and empirical rejection probabilities are small under a fixed coefficient for  $X_2$ , as is obvious from the first row. In Table 2, we also see from rows 2–7 that our test rejects the alternative model more often for a larger variance of  $B_2$ , as we expect. Moreover, the empirical rejection probabilities increase as the covariance of  $B_1$  and  $B_2$  becomes larger, as we see by comparing rows 2–4 with 5–7.

## 5 Application

### 5.1 Motivation: Consumer Demand

Heterogeneity plays an important role in classical consumer demand. The most popular class of parametric demand systems is the almost ideal (AI) class, pioneered by Deaton and Muellbauer [1980]. In the AI model, instead of quantities budget shares are being considered and they are being explained by log prices and log total expenditure<sup>8</sup>. The model is linear in log prices and a term that involves log total expenditure over a nonlinear price index that depends on

<sup>8</sup>The use of total expenditure as wealth concept is standard practice in the demand literature and, assuming the existence of preferences, is satisfied under an assumption of separability of the labor supply from the consumer demand decision, see Lewbel [1999].

parameters of the utility function. In applications, one frequent shortcut is to replace the price index by an actual price index, another is that homogeneity of degree zero is imposed, which means that all prices and total expenditure are relative to a price index, resulting in an entirely linear model.

A popular extension of this model allows for quadratic terms in total expenditure (QUAIDS, Banks et al. [1997]). Since we focus in this paper on the budget share for food at home ( $BSF$ ), which, due at least in parts to satiation effects, is often documented to decline steadily across the total expenditure range, we want to assess whether quadratic terms are really necessary. Note that prices enter the quadratic term in a nonlinear fashion, however, due to the fact that we have very limited price variation, we can treat the nonlinear expression involving prices as fixed. This justifies the use of real total expenditure as regressor, even in the quadratic term. In other words, we thus consider an Engel curve QUAIDS model. However, we want to allow for preference heterogeneity, and hence consider the following model:

$$BSF_i = B_{0i} + B_{1i} \log(TotExp_i) + B_{2i} (\log(TotExp_i))^2 + b_4 W_{1i} + b_5 W_{2i}. \quad (5.1)$$

Unobserved heterogeneity is reflected in the three random coefficients  $B_{0i}$ ,  $B_{1i}$  and  $B_{2i}$ . To account for observed heterogeneity in preferences, we include in addition household covariates as regressors. Specifically, we use principal components to reduce the vector of remaining household characteristics to a few orthogonal, approximately continuous components. We only use two principal components, denoted  $W_{1i}$  and  $W_{2i}$ . While including additional controls in this form is arguably ad hoc, we perform some robustness checks like alternating the component or adding several others, and the results do not change appreciably. Moreover, the additive specification can be justified as letting the mean of the random intercept  $B_{0i}$  depend on covariates.

We implement the test statistics as described in the Monte Carlo section. For testing degeneracy, we estimate the conditional characteristic functions as described in Example 5. For testing functional form restrictions, our test is implemented as described in Example 4, where in the linear case we employ the estimation procedure in Example 7. In both cases, we choose the dimension parameters  $k_n$  and  $m_n$  by the minimum-maximum principle explained in the Monte Carlo section.

## 5.2 The Data: The British Family Expenditure Survey

The FES reports a yearly cross section of labor income, expenditures, demographic composition, and other characteristics of about 7,000 households. We use years 2008 and 2009. As is standard in the demand system literature, we focus on the subpopulation of two person households where both are adults, at least one is working, and the head of household is a white collar worker.



This is to reduce the impact of measurement error; see Lewbel [1999] for a discussion. We thus have a sample of size 543, which is similar to the one considered in the Monte Carlo section.

We form several expenditure categories, but focus on the food at home category. This category contains all food expenditure spent for consumption at home; it is broad since more detailed accounts suffer from infrequent purchases (the recording period is 14 days) and are thus often underreported. Food consumption accounts for roughly 20% of total expenditure. Results actually displayed were generated by considering consumption of food versus nonfood items. We removed outliers by excluding the upper and lower 2.5% of the population in the three groups. We form food budget shares by dividing the expenditures for all food items by total expenditures, as is standard in consumer demand. The following table provides summary statistics of the economically important variables. Since the data are similar to the data used in Hoderlein (2011), for brevity of exposition we refer to this paper for additional descriptive statistics, especially regarding household covariates.

	Min.	1st Qu.	Median	Mean	3rd Qu.	Max.	St. Dev.
Food share	0.008	0.137	0.178	0.188	0.232	0.591	0.075
log(TotExp)	4.207	5.534	5.788	5.782	6.066	6.927	0.448

### 5.3 Results

For testing degeneracy of the coefficient  $B_2$ , we estimate the coefficient under  $H_{\text{deg}}$ , i.e., we assume that this coefficient is fixed. The ordinary least squares estimate is  $-0.009$  with standard error 0.008, which means that mean effects are rather insignificant. A potential role of the nonlinear term more generally is, however, picked up by our procedure. Table 3 shows the different values of the test statistics and p-values at nominal level 0.05. As we see from Table 3, our test fails to reject the model (5.1) with degenerate  $B_{2i}$  but rejects the linear random coefficient model where  $B_{2i} = 0$ . Not surprisingly, we also fail to reject the random coefficient QUAIDS model. The dimension parameters  $k_n$  and  $m_n$  are chosen via the proposed minimum-maximum principle. It is interesting to note that the procedure selects higher order basis functions to account for the random coefficient of the quadratic term. Since higher order basis functions are required to estimate sharp peaks, this also supports the hypothesis that the marginal p.d.f.  $B_2$  is akin to a Dirac measure (i.e., the distribution is degenerate), or very close it.

The analysis thus far assumes that total expenditure is exogenous. However, in consumer demand it is commonly thought that log total expenditure is endogenous and is hence instru-

Null Hypothesis	linear RC	quadratic RC	RC with fixed coeff. on quadratic term in TotExp
value of test	2.1289	1.4200	1.4029
p-values	0.0166	0.0778	0.0803

Table 3: *Values of the tests with p-values when null hypothesis is either a linear random coefficient model (i.e.,  $B_{2i} = 0$  in (5.1)), a quadratic random coefficient model (i.e., random  $B_{2i}$  in (5.1)), or a random coefficient model with degenerate coefficient on the quadratic term (i.e.,  $B_{2i} = b_2$  in (5.1) for some fixed  $b_2$ ).*

mented for, typically by labor income, say  $Z$ , see Lewbel [1999]. One might thus argue that we reject our hypotheses not due to a failure of the functional form assumptions, but because of a violation of exogeneity of total expenditure. Therefore, we follow Imbens and Newey [2009], and model endogeneity through a structural heterogeneous equation that relates total expenditure  $X$  to the instrument labor income  $Z$ , i.e.,

$$X = \psi(Z, U),$$

where  $U$  denotes a scalar unobservable. Following Imbens and Newey [2009], we assume that the instrument  $Z$  is exogenous, i.e., we assume  $Z \perp (B, U)$ , implying  $X \perp B|U$ , and we assume that the function  $\psi$  is strictly monotonic in  $U$ . Finally, we employ the common normalization that  $U|Z$  is uniformly distributed on the unit interval  $[0, 1]$ . Then, the disturbance  $U$  is identified through the conditional cumulative distribution function of  $X$  given  $Z$ , i.e.,

$$U = F_{X|Z}(X|Z).$$

Since  $X \perp B|U$ , we then simply modify our testing procedure by additionally conditioning on controls  $U$ . In the consumer demand literature, this control function approach was also considered by Hoderlein [2011], who propose a life-cycle structural model that yields this specification. Generally, the control function  $U$  would have to be estimated in a first stage. Since the theory involving pre-estimation is beyond the scope of this paper, we do not adjust for estimation error in this variable, which may lead to a higher variance (depending on the smoothness assumptions).

The results of this modification are summarized in Table 4. As we see from this table, the

Null Hypothesis	linear RC	quadratic RC	RC with fixed coeff. on quadratic term in TotExp
value of test	2.0661	1.3978	1.3747
p-values	0.0194	0.0810	0.0846

Table 4: *Values of the test statistics with p-values, when additionally corrected for endogeneity.*

value of the modified test statistics are smaller, once we introduce the instrument  $Z$  in a control function approach. This possibly indicates that there is some endogeneity bias in the first case; however, our main conclusions remain unchanged: We soundly reject the linear RC model, and fail to reject  $H_{\text{deg}}$  and  $H_{\text{mod}}$ .

## 6 Conclusion

This paper develops nonparametric specification testing for random coefficient models. We employ a sieve strategy to obtain tests for both the functional form of the structural equation, e.g., for linearity in random parameters, as well as for the important question of whether or not a parameter can be omitted. While the former can be used to distinguish between various models, including such models where the density of random coefficients is not necessarily point identified, the latter types of test reduce the dimensionality of the random coefficients density. From a nonparametric perspective, this is an important task, because random coefficient models are known to suffer from very slow rates of convergence, see Hoderlein et al. [2010]. We establish the large sample behavior of our test statistics, and show that our tests work well in a finite sample experiment, and allow to obtain reasonable results in a consumer demand application.

## Mathematical Appendix

Throughout the proofs, we will use  $C > 0$  to denote a generic finite constant that may be different in different uses. We use the notation  $a_n \lesssim b_n$  to denote  $a_n \leq Cb_n$  for all  $n \geq 1$ . Further, for ease of notation we write  $\sum_j$  for  $\sum_{j=1}^n$ . Recall that  $\|\cdot\|$  denotes the usual Euclidean norm, while for a matrix  $A$ ,  $\|A\|$  is the operator norm. Further,  $\|\phi\|_X \equiv \sqrt{E(\phi^2(X))}$  and  $\langle \phi, \psi \rangle_X \equiv E[\phi(X)\psi(X)]$ . For any integer  $m \geq 1$ ,  $I_m$  denotes the  $m \times m$  dimensional identity matrix. Recall the notation  $P_n = E[p_{\underline{m}_n}(X)p_{\underline{m}_n}(X)']$ .

### Proofs of Section 2.

PROOF OF PROPOSITION 1. Let us denote  $f = \frac{dF_X}{d\nu}$ . For some constant  $0 < c < 1$ , for all  $n \geq 1$ , and any  $a_n \in \mathbb{R}^{m_n}$  we have

$$\begin{aligned} \|a_n\|^2 &= \int (a_n' p_{\underline{m}_n}(x))^2 1\{f(x) \geq \lambda_n\} \nu(dx) + \int (a_n' p_{\underline{m}_n}(x))^2 1\{f(x) < \lambda_n\} \nu(dx) \\ &\leq \lambda_n^{-1} \int (a_n' p_{\underline{m}_n}(x))^2 f(x) \nu(dx) + c \int (a_n' p_{\underline{m}_n}(x))^2 \nu(dx). \end{aligned}$$

Consequently, we obtain  $\lambda_n I_{m_n} \lesssim P_n$ . □

By Assumption 2, the eigenvalues of  $\lambda_n^{-1}P_n$  are bounded away from zero and hence, it may be assumed that  $P_n = \lambda_n I_{m_n}$ . Otherwise, consider a linear transformation of  $p_{m_n}$  of the form  $\check{p}_{m_n} \equiv (P_n/\lambda_n)^{-1/2}p_{m_n}$  where  $\sup_{x \in \mathcal{X}} \|\check{p}_{m_n}(x)\| \lesssim m_n$  using that the smallest eigenvalue of  $(P_n/\lambda_n)^{-1/2}$  is bounded away from zero uniformly in  $n$ .

**Lemma 6.1.** *It holds  $\sqrt{m_n} \lesssim \varsigma_{m_n}$ .*

*Proof.* Without loss of generality it may be assumed that  $\int \varpi(t)dt = 1$ . By the definition of  $\varsigma_{m_n}$  we conclude

$$\begin{aligned}
\varsigma_{m_n}^2 &\geq \lambda_n^{-2} \sum_{l=1}^{m_n} \int \int \left| E[\delta(V, s) \overline{\delta(V, t)} p_l^2(X)] \right|^2 \varpi(s) \varpi(t) ds dt \\
&\geq \lambda_n^{-2} \sum_{l=1}^{m_n} \left( E[|\int \delta(V, t) \varpi(t) dt|^2 p_l^2(X)] \right)^2 && \text{(by Jensen's inequality)} \\
&\geq C \lambda_n^{-2} \sum_{l=1}^{m_n} (E[p_l^2(X)])^2 && \text{(by Assumption 3)} \\
&= C m_n.
\end{aligned}$$

□

In the following, we make use of the notations  $\widehat{P}_n = n^{-1} \sum_j p_{m_n}(X_j) p_{m_n}(X_j)'$  and  $\widehat{\gamma}_n(t) \equiv (n \widehat{P}_n)^{-1} \sum_j \exp(itY_j) p_{m_n}(X_j)$ . Let  $\widehat{A}_n = n^{-1} \sum_j \int (\mathcal{F}_g q_{k_n})(X_j, t) (\mathcal{F}_g q'_{k_n})(X_j, t) \varpi(t) dt$  and  $A_n = E[\int (\mathcal{F}_g q_{k_n})(X, t) (\mathcal{F}_g q'_{k_n})(X, t) \varpi(t) dt]$ . Recall  $\widehat{\beta}_n = (n \widehat{A}_n)^{-1} \sum_j \int (\mathcal{F}_g q_{k_n})(X_j, t) \widehat{\varphi}_n(X_j, t) \varpi(t) dt$  and let  $\beta_n = A_n^{-1} \int E[(\mathcal{F}_g q_{k_n})(X, t) \varphi(X, t)] \varpi(t) dt$ .

**PROOF OF THEOREM 2.1.** We make use of the decomposition

$$\begin{aligned}
nS_n &= \sum_j \int |\widehat{\varepsilon}_n(X_j, t)|^2 \varpi(t) dt \\
&= \sum_j \int |p_{m_n}(X_j)' \widehat{\gamma}_n(t) - \Pi_{m_n} \varphi(X_j, t)|^2 \varpi(t) dt \\
&\quad + 2 \sum_j \int (p_{m_n}(X_j)' \widehat{\gamma}_n(t) - \Pi_{m_n} \varphi(X_j, t)) (\Pi_{m_n} \varphi(X_j, t) - (\mathcal{F}_g \widehat{f}_{B_n})(X_j, t)) \varpi(t) dt \\
&\quad + \sum_j \int |\Pi_{m_n} \varphi(X_j, t) - (\mathcal{F}_g \widehat{f}_{B_n})(X_j, t)|^2 \varpi(t) dt \\
&= I_n + 2 II_n + III_n \quad (\text{say}).
\end{aligned}$$

Consider  $I_n$ . We conclude

$$\begin{aligned}
I_n &= n \int \left( \widehat{\gamma}_n(t) - \langle \varphi(\cdot, t), p_{\underline{m}_n} \rangle_X \right)' \widehat{P}_n \left( \widehat{\gamma}_n(t) - \langle \varphi(\cdot, t), p_{\underline{m}_n} \rangle_X \right) \varpi(t) dt \\
&= n^{-1} \int \left( \sum_j \left( \exp(itY_j) - \Pi_{m_n} \varphi(X_j, t) \right) p_{\underline{m}_n}(X_j) \right)' \widehat{P}_n^{-1} \\
&\quad \times \left( \sum_j \left( \exp(itY_j) - \Pi_{m_n} \varphi(X_j, t) \right) p_{\underline{m}_n}(X_j) \right) \varpi(t) dt \\
&= \lambda_n^{-1} \int \left\| n^{-1/2} \sum_j \left( \exp(itY_j) - \Pi_{m_n} \varphi(X_j, t) \right) p_{\underline{m}_n}(X_j) \right\|^2 \varpi(t) dt \\
&\quad + n^{-1} \int \left( \sum_j \left( \exp(itY_j) - \Pi_{m_n} \varphi(X_j, t) \right) p_{\underline{m}_n}(X_j) \right)' (\widehat{P}_n^{-1} - \lambda_n^{-1} I_n) \\
&\quad \times \left( \sum_j \left( \exp(itY_j) - \Pi_{m_n} \varphi(X_j, t) \right) p_{\underline{m}_n}(X_j) \right) \varpi(t) dt \\
&= B_{1n} + B_{2n} \quad (\text{say}).
\end{aligned}$$

Since  $(\Pi_{m_n} \varphi(X, t) - \varphi(X, t)) p_{\underline{m}_n}(X)$  is a centered random variable for all  $t$  it is easily seen that  $B_{1n} = \lambda_n^{-1} \int \left\| n^{-1/2} \sum_j \left( \exp(itY_j) - \varphi(X_j, t) \right) p_{\underline{m}_n}(X_j) \right\|^2 \varpi(t) dt + o_p(1)$ . Thus, Lemma 6.2 yields  $(\sqrt{2} \zeta_{m_n})^{-1} (B_{1n} - \mu_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1)$ . To show that  $B_{2n} = o_p(\sqrt{m_n})$  note that

$$\|\widehat{P}_n^{-1} - \lambda_n^{-1} I_n\| \leq \lambda_n^{-1} \|(\widehat{P}_n / \lambda_n)^{-1}\| \|I_n - \widehat{P}_n / \lambda_n\| = O_p\left(\sqrt{(m_n \log n) / (n \lambda_n^2)}\right)$$

by Lemma 6.2 of Belloni et al. [2015]. Further, from  $E[(\exp(itY) - \Pi_{m_n} \varphi(X, t)) p_l(X)] = 0$ ,  $1 \leq l \leq m_n$ , we deduce

$$\begin{aligned}
&n^{-1} \int E \left\| \sum_j \left( \exp(itY_j) - \Pi_{m_n} \varphi(X_j, t) \right) p_{\underline{m}_n}(X_j) \right\|^2 \varpi(t) dt \\
&\lesssim \int \varpi(t) dt E \|p_{\underline{m}_n}(X)\|^2 + \sup_{x \in \mathcal{X}} \|p_{\underline{m}_n}(x)\|^2 \sum_{l=1}^{m_n} \int \langle \varphi(\cdot, t), p_l \rangle_X^2 \varpi(t) dt E[p_l^2(X)] \\
&\lesssim m_n \lambda_n. \tag{6.1}
\end{aligned}$$

The result follows due to condition  $m_n^2 \log n = o(n \lambda_n)$ . Thereby, it is sufficient to prove  $II_n + III_n = o_p(\sqrt{m_n})$ . Consider  $III_n$ . We observe

$$III_n \lesssim \sum_j \int |\mathcal{F}_g(\widehat{f}_{B_n} - \Pi_{k_n} f_B)(X_j, t)|^2 \varpi(t) dt + \sum_j \int |(\mathcal{F}_g \Pi_{k_n} f_B)(X_j, t) - \Pi_{m_n} \varphi(X_j, t)|^2 \varpi(t) dt,$$

where  $\sum_j \int |(\mathcal{F}_g \Pi_{k_n} f_B)(X_j, t) - \Pi_{m_n} \varphi(X_j, t)|^2 \varpi(t) dt = o_p(\sqrt{m_n})$  and

$$\begin{aligned} \sum_j \int |\mathcal{F}_g(\widehat{f}_{B_n} - \Pi_{k_n} f_B)(X_j, t)|^2 \varpi(t) dt &= (\widehat{\beta}_n - \beta_n)' \sum_j \int (\mathcal{F}_g q_{\underline{k}_n})(X_j, t) (\mathcal{F}_g q_{\underline{k}_n})(X_j, t)' \varpi(t) dt (\widehat{\beta}_n - \beta_n) \\ &= n(\widehat{\beta}_n - \beta_n)' \widehat{A}_n (\widehat{\beta}_n - \beta_n). \end{aligned}$$

Let us introduce the vector  $\widetilde{\beta}_n = (n\widehat{A}_n)^- \sum_j \int (\mathcal{F}_g q_{\underline{k}_n})(X_j, t) \varphi(X_j, t) \varpi(t) dt$ . Using the property of Moore-Penrose inverses that  $\widehat{A}_n = \widehat{A}_n \widehat{A}_n^- \widehat{A}_n$ , we conclude

$$\begin{aligned} n(\widehat{\beta}_n - \beta_n)' \widehat{A}_n (\widehat{\beta}_n - \beta_n) &\lesssim n(\widehat{\beta}_n - \widetilde{\beta}_n)' \widehat{A}_n (\widehat{\beta}_n - \widetilde{\beta}_n) + n(\widetilde{\beta}_n - \beta_n)' \widehat{A}_n (\widetilde{\beta}_n - \beta_n) \\ &\lesssim \left\| n^{-1/2} \sum_j \int (\mathcal{F}_g q_{\underline{k}_n})(X_j, t) (\widehat{\varphi}_n(X_j, t) - \varphi(X_j, t)) \varpi(t) dt \right\|^2 \|\widehat{A}_n^-\| \\ &+ \left\| n^{-1/2} \sum_j \int (\mathcal{F}_g q_{\underline{k}_n})(X_j, t) \varphi(X_j, t) \varpi(t) dt \right\|^2 \|\widehat{A}_n^- - A_n^-\|^2 \|\widehat{A}_n\| \\ &+ \left\| n^{-1/2} \sum_j \int \left( (\mathcal{F}_g q_{\underline{k}_n})(X_j, t) \varphi(X_j, t) - E[(\mathcal{F}_g q_{\underline{k}_n})(X, t) \varphi(X, t)] \right) \varpi(t) dt \right\|^2 \|A_n^-\|^2 \|\widehat{A}_n\|. \end{aligned}$$

From Lemma 6.3 we have  $\|\widehat{A}_n^- - A_n^-\| = O_p(\sqrt{(\log n)k_n/n})$ . By Assumption 4 (v) it holds  $\|A_n^-\| = O(1)$  and thus,  $\|\widehat{A}_n^-\| \leq \|\widehat{A}_n^- - A_n^-\| + \|A_n^-\| = O_p(1)$ . Thereby, it is sufficient to consider

$$\begin{aligned} &\left\| n^{-1/2} \sum_j \int (\mathcal{F}_g q_{\underline{k}_n})(X_j, t) (\widehat{\varphi}_n(X_j, t) - \varphi(X_j, t)) \varpi(t) dt \right\|^2 \\ &\lesssim \left\| n^{-1/2} \sum_j \int (\mathcal{F}_g q_{\underline{k}_n})(X_j, t) p_{\underline{m}_n}(X_j)' (\widehat{\gamma}_n(t) - \langle \varphi(\cdot, t), p_{\underline{m}_n} \rangle_X) \varpi(t) dt \right\|^2 \\ &\quad + \left\| n^{-1/2} \sum_j \int (\mathcal{F}_g q_{\underline{k}_n})(X_j, t) (\Pi_{m_n} \varphi(X_j, t) - \varphi(X_j, t)) \varpi(t) dt \right\|^2 \\ &\lesssim n \left\| \int E[(\mathcal{F}_g q_{\underline{k}_n})(X, t) p_{\underline{m}_n}(X)'] (\widehat{\gamma}_n(t) - \langle \varphi(\cdot, t), p_{\underline{m}_n} \rangle_X) \varpi(t) dt \right\|^2 \\ &\quad + n \left\| \int E[(\mathcal{F}_g q_{\underline{k}_n})(X, t) (\Pi_{m_n} \varphi(X, t) - \varphi(X, t))] \varpi(t) dt \right\|^2 + O_p(k_n) \\ &= O_p(k_n + n \|\Pi_{m_n} \varphi - \varphi\|_{\varpi}^2) \end{aligned}$$

which can be seen as follows. Let  $\langle \cdot, \cdot \rangle_{\varpi}$  denote the inner product induced by the norm  $\|\cdot\|_{\varpi}$ .

We calculate

$$\begin{aligned}
\left\| \int E[(\mathcal{F}_g q_{\underline{k}_n})(X, t)(\Pi_{m_n} \varphi(X, t) - \varphi(X, t))] \varpi(t) dt \right\|^2 &= \sum_{l=1}^{k_n} \langle \mathcal{F}_g q_l, \Pi_{m_n} \varphi - \varphi \rangle_{\varpi}^2 \\
&= \sum_{l=1}^{k_n} \left( \int q_l(b) E[(\mathcal{F}_g^*(\Pi_{m_n} \varphi - \varphi))(X, b)] db \right)^2 \\
&\lesssim \int \left( E[(\mathcal{F}_g^*(\Pi_{m_n} \varphi - \varphi))(X, b)] \right)^2 db \\
&\lesssim \|\Pi_{m_n} \varphi - \varphi\|_{\varpi}^2
\end{aligned}$$

where  $\mathcal{F}_g^*$  is the adjoint operator of  $\mathcal{F}_g$  given by  $(\mathcal{F}_g^* \phi)(b) = \int E[\exp(itg(X, b)) \phi(X, t)] \varpi(t) dt$ . Consequently, we have  $n(\widehat{\beta}_n - \beta_n)' \widehat{A}_n (\widehat{\beta}_n - \beta_n) = O_p(k_n + n \|\Pi_{m_n} \varphi - \varphi\|_{\varpi}^2) = o_p(\sqrt{m_n})$  and, in particular,  $III_n = o_p(\sqrt{m_n})$ . Consider  $II_n$ . From above we infer  $n \|\widehat{\beta}_n - \beta_n\|^2 = O_p(k_n + n \|\Pi_{m_n} \varphi - \varphi\|_{\varpi}^2)$  by employing that  $\widehat{A}_n^-$  is stochastically bounded. Thereby, we obtain

$$\begin{aligned}
|II_n|^2 &\lesssim \left| \int \sum_j (\exp(itY_j) - \Pi_{m_n} \varphi(X_j, t)) p_{\underline{m}_n}(X_j)' \langle \varphi(\cdot, t) - (\mathcal{F}_g \Pi_{k_n} f_B)(\cdot, t), p_{\underline{m}_n} \rangle_X \varpi(t) dt \right|^2 \\
&\quad + \left| \int \sum_j (\exp(itY_j) - \Pi_{m_n} \varphi(X_j, t)) p_{\underline{m}_n}(X_j)' \langle (\mathcal{F}_g(\Pi_{k_n} f_B - \widehat{f}_{B_n}))(\cdot, t), p_{\underline{m}_n} \rangle_X \varpi(t) dt \right|^2 + o_p(m_n) \\
&\lesssim n \int E |\Pi_{m_n} \varphi(X, t) - (\Pi_{m_n} \mathcal{F}_g \Pi_{k_n} f_B)(X, t)|^2 \varpi(t) dt \\
&\quad + n \int E \left\| (\exp(itY) - \Pi_{m_n} \varphi(X, t)) (\Pi_{m_n} \mathcal{F}_g q_{\underline{k}_n})(X, t) \right\|^2 \varpi(t) dt \|\widehat{\beta}_n - \beta_n\|^2 + o_p(m_n) \\
&= O_p \left( n \|\mathcal{F}_g(\Pi_{k_n} f_B - f_B)\|_{\varpi}^2 + k_n (k_n + n \|\Pi_{m_n} \varphi - \varphi\|_{\varpi}^2) \right) + o_p(m_n) \\
&= o_p(m_n)
\end{aligned}$$

where we used that  $\|\Pi_{m_n} \varphi - \varphi\|_{\infty} = O(1)$  and  $\sum_{l=1}^{k_n} \|\Pi_{m_n} \mathcal{F}_g q_l\|_{\varpi}^2 = O(k_n)$ , which completes the proof.  $\square$

We require the following notation. Let us introduce the covariance matrix estimator  $\widehat{\Sigma}_{m_n}(s, t) = n^{-1} \sum_j p_{\underline{m}_n}(X_j) p_{\underline{m}_n}(X_j)' \delta_n(V_j, s) \delta_n(V_j, t)$  where  $\delta_n(V_j, s) = \exp(itY) - (\mathcal{F}_g \widehat{f}_{B_n})(X, t)$ . Further, we define  $\widetilde{\delta}_n(V, t) = \exp(itY) - (\mathcal{F}_g \Pi_{k_n} f_B)(X, t)$  and introduce the matrix  $\widetilde{\Sigma}_{m_n}(s, t) = n^{-1} \sum_j p_{\underline{m}_n}(X_j) p_{\underline{m}_n}(X_j)' \widetilde{\delta}_n(V_j, s) \widetilde{\delta}_n(V_j, t)$ .

**PROOF OF PROPOSITION 2.** To keep the presentation of the proof simple, we do not consider

estimation of  $P_n$  in  $\widehat{\Sigma}_{m_n}$  and  $\widehat{\mu}_{m_n}$ . We make use of the relationship

$$\begin{aligned} \delta_n(\cdot, s)\delta_n(\cdot, t) - \widetilde{\delta}_n(\cdot, s)\widetilde{\delta}_n(\cdot, t) &= \widetilde{\delta}_n(\cdot, s)((\mathcal{F}_g\widehat{f}_{B_n})(\cdot, t) - (\mathcal{F}_g\Pi_{k_n}f_B)(\cdot, t)) \\ &\quad + \delta_n(\cdot, t)((\mathcal{F}_g\widehat{f}_{B_n})(\cdot, s) - (\mathcal{F}_g\Pi_{k_n}f_B)(\cdot, s)). \end{aligned}$$

Observe

$$\begin{aligned} &\int \int \|\widehat{\Sigma}_{m_n}(s, t) - \widetilde{\Sigma}_{m_n}(s, t)\|_F^2 \varpi(s)ds\varpi(t)dt \\ &\lesssim \int \int \left\| n^{-1} \sum_j p_{\underline{m}_n}(X_j)p_{\underline{m}_n}(X_j)' \widetilde{\delta}_n(V_j, s)((\mathcal{F}_g\widehat{f}_{B_n})(X_j, t) - (\mathcal{F}_g\Pi_{k_n}f_B)(X_j, t)) \right\|_F^2 \varpi(s)ds\varpi(t)dt \\ &+ \int \int \left\| n^{-1} \sum_j p_{\underline{m}_n}(X_j)p_{\underline{m}_n}(X_j)' \delta_n(V_j, t)((\mathcal{F}_g\widehat{f}_{B_n})(X_j, s) - (\mathcal{F}_g\Pi_{k_n}f_B)(X_j, s)) \right\|_F^2 \varpi(s)ds\varpi(t)dt \\ &= I_n + II_n \quad (\text{say}). \end{aligned}$$

We conclude

$$\begin{aligned} I_n &\leq \int \int \left\| \frac{1}{n} \sum_j \widetilde{\delta}_n(V_j, s)p_{\underline{m}_n}(X_j)p_{\underline{m}_n}(X_j)'(\mathcal{F}_g p_{\underline{k}_n})(X_j, t)'(\widehat{\beta}_n - \beta_n) \right\|_F^2 \varpi(s)ds\varpi(t)dt \\ &\leq \int \int \left\| E[\widetilde{\delta}_n(V, s)p_{\underline{m}_n}(X)p_{\underline{m}_n}(X)'(\mathcal{F}_g q_{\underline{k}_n})(X, t)'](\widehat{\beta}_n - \beta_n) \right\|_F^2 \varpi(s)ds\varpi(t)dt + o_p(1) \\ &\leq \|\widehat{\beta}_n - \beta_n\|^2 \\ &\quad \times O_p\left(\sum_{j,l=1}^{m_n} \sum_{l'=1}^{k_n} \int \int \left(E[(\varphi(X, s) - (\mathcal{F}_g\Pi_{k_n}f_B)(X, s))(\mathcal{F}_g q_{l'})(X, t)p_j(X)p_l(X)]\right)^2 \varpi(s)ds\varpi(t)dt\right) \\ &\leq \|\widehat{\beta}_n - \beta_n\|^2 \\ &\quad \times O_p\left(\sum_{l=1}^{m_n} \sum_{l'=1}^{k_n} \int \int E|(\varphi(X, s) - (\mathcal{F}_g\Pi_{k_n}f_B)(X, s))(\mathcal{F}_g q_{l'})(X, t)p_l(X)|^2 \varpi(s)ds\varpi(t)dt\right) \\ &= O_p\left(m_n k_n^2/n + m_n k_n \|\Pi_{m_n}\varphi - \varphi\|_{\varpi}^2\right) = o_p(1). \end{aligned}$$

Here, we used  $\|\widehat{\beta}_n - \beta_n\|^2 = O_p(k_n/n + \|\Pi_{m_n}\varphi - \varphi\|_{\varpi}^2)$  which can be seen as in the proof of



Theorem 2.1. Since  $I_n = o_p(1)$  we conclude

$$\begin{aligned}
II_n &\lesssim \int \int \left\| (\widehat{\beta}_n - \beta_n)' E[(\mathcal{F}_g q_{k_n})(X, s) p_{m_n}(X) p_{m_n}(X)' (\mathcal{F}_g q_{k_n})(X, t)] (\widehat{\beta}_n - \beta_n) \right\|_F^2 \varpi(s) ds \varpi(t) dt + o_p(1) \\
&\leq \|\widehat{\beta}_n - \beta_n\|^4 \sum_{j,l=1}^{m_n} \int \int E[\|(\mathcal{F}_g q_{k_n})(X, s)\| \|(\mathcal{F}_g q_{k_n})(X, t)\| |p_j(X) p_l(X)|^2] \varpi(s) ds \varpi(t) dt + o_p(1) \\
&\leq C m_n^2 \|\widehat{\beta}_n - \beta_n\|^4 \left( \int E\|(\mathcal{F}_g q_{k_n})(X, t)\|^2 \varpi(t) dt \right)^2 + o_p(1) \\
&= O_p(m_n^2 k_n / n^2 + m_n^2 \|\Pi_{m_n} \varphi - \varphi\|_{\varpi}^4) = o_p(1).
\end{aligned}$$

by using  $k_n = o(\sqrt{m_n})$ . Finally, it is easily seen that  $\varsigma_{m_n}^2 - \int \int \|\widetilde{\Sigma}_{m_n}(s, t)\|^2 \varpi(s) ds \varpi(t) dt = o_p(1)$ , which proves  $\varsigma_{m_n} \widehat{\varsigma}_{m_n}^{-1} = 1 + o_p(1)$ . In particular, convergence of the trace of  $\widehat{\Sigma}_{m_n}(t, t)$  to the trace of  $\Sigma_{m_n}(t, t)$  follows by using  $|\widehat{\mu}_{m_n} - \mu_{m_n}|^2 \leq m_n \int \|\widehat{\Sigma}_{m_n}(t, t) - \Sigma_{m_n}(t, t)\|_F^2 \varpi(t) dt = o_p(m_n)$ .  $\square$

PROOF OF THEOREM 2.2. Let us introduce  $\alpha_n = (n\mathcal{A}_n)^{-} \int E[\chi_n(X, t) \varphi(X, t)] \varpi(t) dt$  and the estimator

$$\widehat{\alpha}_n = (n\widehat{\mathcal{A}}_n)^{-} \int \sum_j \chi_n(X_j, t) \widehat{\varphi}_n(X_j, t) \varpi(t) dt.$$

We prove in the following that

$$\begin{aligned}
&\sum_j \int |\widehat{\varphi}_n(X_j, t) - (\mathcal{F}_{g_1} \widehat{f}_{B_{-2}, n})(X_j, t) \widetilde{g}(X_j, t, \widehat{b}_{2n})|^2 \varpi(t) dt \\
&= \sum_j \int |\widehat{\varphi}_n(X_j, t) - \chi_n(X_j, t) \widehat{\alpha}_n|^2 \varpi(t) dt + o_p(\sqrt{m_n}).
\end{aligned}$$

By the definition of the estimator  $\widehat{b}_{2n}$  in (2.8) we obtain

$$\begin{aligned}
&\sum_j \int |\widehat{\varphi}_n(X_j, t) - (\mathcal{F}_{g_1} \widehat{f}_{B_{-2}, n})(X_j, t) \widetilde{g}(X_j, t, \widehat{b}_{2n})|^2 \varpi(t) dt \\
&\leq \sum_j \int |\widehat{\varphi}_n(X_j, t) - (\mathcal{F}_{g_1} \widehat{f}_{B_{-2}, n})(X_j, t) \widetilde{g}(X_j, t, b_2)|^2 \varpi(t) dt \quad (6.2)
\end{aligned}$$

for any  $b_2 \in \mathcal{B}_2$  satisfying (2.5). By the definition of the least squares estimator  $\widehat{\alpha}_n$  and the

triangular inequality we obtain

$$\begin{aligned}
& \sqrt{\sum_j \int |\widehat{\varphi}_n(X_j, t) - (\mathcal{F}_{g_1} \widehat{f}_{B_{-2}, n})(X_j, t) \widetilde{g}(X_j, t, \widehat{b}_{2n})|^2 \varpi(t) dt} \\
& \geq \sqrt{\sum_j \int |\widehat{\varphi}_n(X_j, t) - (\mathcal{F}_{g_1} \widehat{f}_{B_{-2}, n})(X_j, t) \Pi_{l_n} \widetilde{g}(X_j, t, \widehat{b}_{2n})|^2 \varpi(t) dt} \\
& \quad - \sqrt{\sum_j \int |(\mathcal{F}_{g_1} \widehat{f}_{B_{-2}, n})(X_j, t) (\Pi_{l_n} \widetilde{g}(X_j, t, \widehat{b}_{2n}) - \widetilde{g}(X_j, t, \widehat{b}_{2n}))|^2 \varpi(t) dt} \\
& \geq \sum_j \int |\widehat{\varphi}_n(X_j, t) - \chi_n(X_j, t)' \widehat{\alpha}_n|^2 \varpi(t) dt - O_p\left(\sqrt{n} \max_{b \in \mathcal{B}_2} \|\Pi_{l_n} \widetilde{g}(\cdot, \cdot, b) - \widetilde{g}(\cdot, \cdot, b)\|_{\varpi}\right) \\
& = \sum_j \int |\widehat{\varphi}_n(X_j, t) - \chi_n(X_j, t)' \widehat{\alpha}_n|^2 \varpi(t) dt - o_p(m_n^{1/4}).
\end{aligned}$$

Consequently, applying again the triangular inequality together with inequality (6.2) yields

$$\begin{aligned}
& \left| \sqrt{\sum_j \int |\widehat{\varphi}_n(X_j, t) - (\mathcal{F}_{g_1} \widehat{f}_{B_{-2}, n})(X_j, t) \widetilde{g}(X_j, t, \widehat{b}_{2n})|^2 \varpi(t) dt} \right. \\
& \quad \left. - \sqrt{\sum_j \int |\widehat{\varphi}_n(X_j, t) - \chi_n(X_j, t)' \widehat{\alpha}_n|^2 \varpi(t) dt} \right| \\
& \leq \sqrt{\sum_j \int |\widehat{\varphi}_n(X_j, t) - (\mathcal{F}_{g_1} \widehat{f}_{B_{-2}, n})(X_j, t) \widetilde{g}(X_j, t, b_2)|^2 \varpi(t) dt} \\
& \quad - \sqrt{\sum_j \int |\widehat{\varphi}_n(X_j, t) - \chi_n(X_j, t)' \widehat{\alpha}_n|^2 \varpi(t) dt} + o_p(m_n^{1/4}) \\
& \leq \sqrt{\sum_j \int |\Pi_{l_n} \widetilde{g}(X_j, t, b_2) - \widetilde{g}(X_j, t, b_2)|^2 \varpi(t) dt} + \sqrt{n} \|\widehat{\alpha}_n - \alpha_n\| + o_p(m_n^{1/4}) \\
& = n \|\Pi_{l_n} \widetilde{g}(\cdot, \cdot, b_2) - \widetilde{g}(\cdot, \cdot, b_2)\|_{\varpi}^2 + O_p\left(\sqrt{k_n l_n} + \sqrt{n} \|\Pi_{m_n} \varphi - \varphi\|_{\varpi}\right) + o_p(m_n^{1/4}) \\
& = o_p(m_n^{1/4}),
\end{aligned}$$

as in the proof of Theorem 2.1. Now following line by line the proof of Theorem 2.1 and using

$$\begin{aligned}
& \sum_j \int |\Pi_{m_n} \varphi(X_j, t) - (\mathcal{F}_{g_1} \Pi_{k_n} f_{B_{-2}})(X_j, t) \Pi_{l_n} \widetilde{g}(X_j, t, b_2)|^2 \varpi(t) dt \\
& \lesssim n \|\Pi_{m_n} \varphi - \varphi\|_{\varpi}^2 + n \|\mathcal{F}_{g_1} \Pi_{k_n} f_{B_{-2}} - \mathcal{F}_{g_1} f_{B_{-2}}\|_{\varpi}^2 + n \|\Pi_{l_n} \widetilde{g}(\cdot, \cdot, b_2) - \widetilde{g}(\cdot, \cdot, b_2)\|_{\varpi}^2 + o_p(\sqrt{m_n}) \\
& = o_p(\sqrt{m_n}),
\end{aligned}$$

the result follows.  $\square$

PROOF OF THEOREM 2.3. We make use of the decomposition

$$\begin{aligned}
nS_n &= \sum_j \int |p_{\underline{m}_n}(X_j)'(\widehat{\gamma}_n(t) - \langle \varphi(\cdot, t), p_{\underline{m}_n} \rangle_X)|^2 \varpi(t) dt \\
&\quad + 2 \sum_j \int \left( p_{\underline{m}_n}(X_j)'(\widehat{\gamma}_n(t) - \langle \varphi(\cdot, t), p_{\underline{m}_n} \rangle_X) \right) \\
&\quad \quad \times \left( \Pi_{m_n} \varphi(X_j, t) - \widehat{h}_n(X_{1j}, t) \widetilde{g}(X_j, t, \widehat{b}_{2n}) \right) \varpi(t) dt \\
&\quad + \sum_j \int |\Pi_{m_n} \varphi(X_j, t) - \widehat{h}_n(X_{1j}, t) \widetilde{g}(X_j, t, \widehat{b}_{2n})|^2 \varpi(t) dt \\
&= I_n + 2 II_n + III_n \quad (\text{say})
\end{aligned}$$

where we used  $\langle h(\cdot, t) \widetilde{g}(\cdot, t, b_2), p_{\underline{m}_n} \rangle_X = \langle \varphi(\cdot, t), p_{\underline{m}_n} \rangle_X$ . Consider  $I_n$ . As in the proof of Theorem 2.1 we have

$$I_n = n\lambda_n^{-1} \int \left\| n^{-1/2} \sum_j \left( \exp(itY_j) - h(X_{1j}, t) \widetilde{g}(X_j, t, b_2) \right) p_{\underline{m}_n}(X_j) \right\|^2 \varpi(t) dt + o_p(\sqrt{m_n}).$$

Thus, Lemma 6.2 yields  $(\sqrt{2}\varsigma_{m_n})^{-1}(I_n - \mu_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1)$ . Consider  $III_n$ . Since  $|\widetilde{g}(X_j, t, b)| \leq 1$  for all  $b$  we evaluate

$$\begin{aligned}
III_n &\lesssim \sum_j \int |\Pi_{m_n} \varphi(X_j, t) - \varphi(X_j, t)|^2 \varpi(t) dt \\
&\quad + \sum_j \int |h(X_{1j}, t) - \widehat{h}_n(X_{1j}, t)|^2 \varpi(t) dt \\
&\quad + \sum_j \int |\widehat{h}_n(X_{1j}, t)|^2 |\widetilde{g}(X_j, t, b_2) - \widetilde{g}(X_j, t, \widehat{b}_{2n})|^2 \varpi(t) dt.
\end{aligned}$$

It holds  $\int \|\widehat{h}_n(\cdot, t) - \Pi_{k_n} h(\cdot, t)\|_{X_1}^2 \varpi(t) dt = O_p(k_n/n)$  as we see in the following. We have

$$\begin{aligned}
&\lambda_n \int \|\widehat{h}_n(\cdot, t) - \Pi_{k_n} h(\cdot, t)\|_{X_1}^2 \varpi(t) dt \\
&\leq \lambda_n \left\| \left( \sum_j p_{\underline{k}_n}(X_j) p_{\underline{k}_n}(X_j)' \right)^{-1} \right\| \int \left\| \sum_j \left( \Pi_{k_n} h(X_{1j}, t) - \exp(it(Y_j - g_2(X_j, \widehat{b}_{2n}))) \right) p_{\underline{k}_n}(X_{1j}) \right\|^2 \varpi(t) dt \\
&\quad \lesssim \int \left\| n^{-1} \sum_j \left( \Pi_{k_n} h(X_{1j}, t) - \exp(it(Y_j - g_2(X_j, b_2))) \right) p_{\underline{k}_n}(X_{1j}) \right\|^2 \varpi(t) dt \\
&\quad + \|\widehat{b}_{2n} - b_2\|^2 \sum_{l=1}^{k_n} \int \left\| n^{-1} \sum_j \exp(itY_j) \widetilde{g}_b(X_j, t, \widetilde{b}_{2n}) p_l(X_{1j}) \right\|^2 \varpi(t) dt + o_p(1),
\end{aligned}$$

by a Taylor series expansion, where  $\widetilde{b}_{2n}$  is between  $\widehat{b}_{2n}$  and  $b_2$ . As in relation (6.1), from

$E[(\Pi_{k_n} h(X, t) - \exp(it(Y - g_2(X, b_2)))p_{\underline{k}_n}(X))] = 0$  we deduce

$$\int E \left\| n^{-1} \sum_j (\Pi_{k_n} h(X_{1j}, t) - \exp(it(Y_j - g_2(X_j, b_2)))p_{\underline{k}_n}(X_j)) \right\|^2 \varpi(t) dt = O(n^{-1} k_n \lambda_n).$$

Further, since  $\int E \sup_{b \in \mathcal{B}_2} \|\tilde{g}_b(X, t, b)\|^2 \varpi(t) dt \leq C$  we have

$$\begin{aligned} E \left( \sum_{l=1}^{k_n} \int \left\| n^{-1} \sum_j \exp(itY_j) \tilde{g}_b(X_j, t, \tilde{b}_{2n}) p_l(X_j) \right\|^2 \varpi(t) dt \right)^{1/2} \\ \leq E \left[ \|p_{\underline{k}_n}(X)\| \left( \int \|\tilde{g}_b(X, t, \tilde{b}_{2n})\|^2 \varpi(t) dt \right)^{1/2} \right] \\ \leq (E \|p_{\underline{k}_n}(X)\|^2)^{1/2} \left( \int E \sup_{b \in \mathcal{B}_2} \|\tilde{g}_b(X, t, b)\|^2 \varpi(t) dt \right)^{1/2} \\ = O(\sqrt{\lambda_n k_n}). \end{aligned}$$

This establishes the rate for the estimator  $\hat{h}_n$ . In light of condition  $n \|\Pi_{k_n} h - h\|_{\varpi}^2 = o(\sqrt{m_n})$ , from  $n \|b_2 - \hat{b}_{2n}\|^2 = O_p(1)$  and  $k_n = o(\sqrt{m_n})$  we infer  $III_n = o_p(\sqrt{m_n})$ . It remains to show  $II_n = o_p(\sqrt{m_n})$ , which follows by

$$\begin{aligned} |II_n| &\lesssim \left| \int \sum_j (\exp(itY_j) - \Pi_{m_n} \varphi(X_j, t)) p_{\underline{m}_n}(X_j)' \langle \Pi_{m_n} \varphi(\cdot, t) - \Pi_{k_n} h(\cdot, t) \tilde{g}(\cdot, t, b_2), p_{\underline{m}_n} \rangle_X \varpi(t) dt \right| \\ &+ \left| \int \sum_j (\exp(itY_j) - \Pi_{m_n} \varphi(X_j, t)) p_{\underline{m}_n}(X_j)' \langle \Pi_{k_n} h(\cdot, t) \tilde{g}(\cdot, t, b_2) - \hat{h}_n(\cdot, t) \tilde{g}(\cdot, t, \hat{b}_{2n}), p_{\underline{m}_n} \rangle_X \varpi(t) dt \right| \\ &+ o_p(\sqrt{m_n}) \\ &= O_p(\sqrt{n} \|\Pi_{k_n} h - h\|_{\varpi}) + o_p(\sqrt{m_n}) \\ &+ O_p \left( \left( k_n \int E \sup_{b \in \mathcal{B}_2} \left\| \sum_{l=1}^{m_n} p_l(X) \langle \tilde{g}_b(\cdot, t, b) p'_{\underline{k}_n}, p_l \rangle_X \right\|^2 \varpi(t) dt \right)^{1/2} \right) \\ &= o_p(\sqrt{m_n}), \end{aligned}$$

using that  $\int E \sup_{b \in \mathcal{B}_2} \left\| \sum_{l=1}^{m_n} p_l(X) \langle \tilde{g}_b(\cdot, t, b) p'_{\underline{k}_n}, p_l \rangle_X \right\|^2 \varpi(t) dt \leq \sum_{l=1}^{k_n} E[p_l^2(X)] = O(k_n)$ , which proves the result.  $\square$

In the following, recall the definition of  $f_B^*$  satisfying  $\|\mathcal{F}_g f_B^* - \varphi\|_{\varpi} \leq \|\mathcal{F}_g f - \varphi\|_{\varpi}$  for all p.d.f.  $f$ .

**PROOF OF PROPOSITION 3.** For the proof it is sufficient to show  $S_n \geq C \|\mathcal{F}_g f_B^* - \varphi\|_{\varpi}^2 + o_p(1)$ .

The proof of Theorem 2.1 together with the basic inequality  $(a - b)^2 \geq a^2 - b^2$  implies that

$$\begin{aligned}
S_n &= \lambda_n^{-1} \sum_{l=1}^{m_n} \int \left| n^{-1} \sum_j (\exp(itY_j) - (\mathcal{F}_g f_B^*)(X_j, t)) p_l(X_j) \right|^2 \varpi(t) dt + o_p(1) \\
&\gtrsim \sum_{l=1}^{m_n} \int |E[(\exp(itY) - (\mathcal{F}_g f_B^*)(X, t)) p_l(X)]|^2 \varpi(t) dt + o_p(1) \\
&\gtrsim \|\mathcal{F}_g f_B^* - \varphi\|_{\varpi}^2 + o_p(1),
\end{aligned}$$

by using that  $(\lambda_n)_{n \geq 1}$  is a nonincreasing sequence. □

PROOF OF PROPOSITION 4. Following the proof Theorem 2.1, it is easily seen that

$$\begin{aligned}
nS_n &= \lambda_n^{-1} \sum_{l=1}^{m_n} \int \left| n^{-1/2} \sum_j (Y_j - \varphi(X_j, t)) p_l(X_j) \right|^2 \varpi(t) dt \\
&\quad + \sum_j \int |(\mathcal{F}_g \Pi_{k_n} f_B^*)(X_j, t) - \Pi_{m_n} \varphi(X_j, t)|^2 \varpi(t) dt + o_p(\sqrt{m_n}).
\end{aligned}$$

Further, under the sequence of local alternatives (2.11), we calculate

$$\begin{aligned}
\sum_j \int |(\mathcal{F}_g \Pi_{k_n} f_B^*)(X_j, t) - \Pi_{m_n} \varphi(X_j, t)|^2 \varpi(t) dt &= n \|\mathcal{F}_g f_B^* - \varphi\|_{\varpi}^2 + o_p(\sqrt{m_n}) \\
&= \varsigma_{m_n}^{-1} \|\mathcal{F}_g \Delta\|_{\varpi}^2 + o_p(\sqrt{m_n}),
\end{aligned}$$

which proves the result. □

### Proofs of Section 3.

In the following, we make use of the notation  $\hat{\alpha}_n \equiv (n\hat{R}_n)^{-1} \sum_j Y_j p_{m_n}(X_j, Z_j)$  where  $\hat{R}_n = n^{-1} \sum_j p_{m_n}(X_j, Z_j) p_{m_n}(X_j, Z_j)'$ . The Kronecker product for matrices is denoted by  $\otimes$ .

PROOF OF PROPOSITION 5. We make use of the decomposition

$$\begin{aligned}
nS_n &= \sum_j \int \left| \partial_z p_{\underline{m}_n}(X_j, z)' (\hat{\alpha}_n - E[1\{Z > g(X, B)\} p_{\underline{m}_n}(X, Z)] \right|^2 \varpi(z) dz \\
&\quad + 2 \sum_j \int \partial_z p_{\underline{m}_n}(X_j, z)' \left( \hat{\alpha}_n - E[1\{Z > g(X, B)\} p_{\underline{m}_n}(X, Z)] \right) \\
&\quad \quad \times \left( \partial_z p_{\underline{m}_n}(X_j, z)' E[1\{Z > g(X, B)\} p_{\underline{m}_n}(X, Z)] - (\mathcal{F}^{-1}[(\mathcal{F}_g \hat{f}_{B_n})(X_j, \cdot)])(z) \right) \varpi(z) dz \\
&\quad + \sum_j \int \left| \partial_z p_{\underline{m}_n}(X_j, z)' E[1\{Z > g(X, B)\} p_{\underline{m}_n}(X, Z)] - (\mathcal{F}^{-1}[(\mathcal{F}_g \hat{f}_{B_n})(X_j, \cdot)])(z) \right|^2 \varpi(z) dz \\
&= I_n + 2II_n + III_n \quad (\text{say}).
\end{aligned}$$

Consider  $I_n$ . For all  $l \geq 1$ , the derivative of a basis function  $p_l$  is given by  $lp_{l-1}$ . Since  $p_l$  forms an orthonormal basis in  $L^2_{\varpi}(\mathbb{R})$  is holds

$$\begin{aligned}
I_n &= \frac{n}{\lambda_n} (\hat{\beta}_{m_n} - E[1\{Z > g(X, B)\} p_{\underline{m}_n}(X, Z)])' (I_{m_{2n}} \otimes T_n) (\hat{\beta}_{m_n} - E[1\{Z > g(X, B)\} p_{\underline{m}_n}(X, Z)]) \\
&\quad + o_p(\sqrt{m_n})
\end{aligned}$$

where  $T_n$  is a  $m_{2n} \times m_{2n}$  diagonal matrix with  $l$ -th diagonal element is given by  $(l-1)^2$ . It holds

$$I_n = \lambda_n^{-1} \sum_{l=1}^{m_n} \tau_l \left| n^{-1/2} \sum_j \left( Y_j - \int 1\{Z_j > g_2(X_j, b)\} f_B(b) db \right) p_l(X_j, Z_j) \right|^2 + o_p(1).$$

Thus, Lemma 6.2 yields  $(\sqrt{2}\varsigma_{m_n})^{-1}(I_n - \mu_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1)$ . Consider  $III_n$ . We have

$$\begin{aligned}
III_n &\lesssim \sum_j \int \left| (\Pi_{m_n} \psi)(X_j, z) - \psi(X_j, z) \right|^2 \varpi(z) dz + \sum_j \int \left| (\mathcal{F}^{-1}[(\mathcal{F}_g(\hat{f}_{B_n} - f_B))(X_j, \cdot)])(z) \right|^2 \varpi(z) dz \\
&= A_{n1} + A_{n2}.
\end{aligned}$$

We have  $A_{n1} = O_p(n \|\Pi_{m_n} \psi - \psi\|_{\varpi}^2) = o_p(\sqrt{m_n})$  and

$$\begin{aligned}
A_{n2} &\lesssim \sum_j \int \left| (\mathcal{F}^{-1}[(\mathcal{F}_g(\hat{f}_{B_n} - \Pi_{k_n} f_B))(X_j, \cdot)])(z) \right|^2 \varpi(z) dz \\
&\quad + \sum_j \int \left| (\mathcal{F}^{-1}[(\mathcal{F}_g(\Pi_{k_n} f_B - f_B))(X_j, \cdot)])(z) \right|^2 \varpi(z) dz,
\end{aligned}$$

where the second summand on the right hand is of the order  $o_p(\sqrt{m_n})$ . Further,

$$\begin{aligned} & \sum_j \int |(\mathcal{F}^{-1}[(\mathcal{F}_g(\widehat{f}_{B_n} - \Pi_{k_n} f_B))(X_j, \cdot)])(z)|^2 \varpi(z) dz \\ &= (\widehat{\beta}_n - \beta_n)' \sum_j \int (\mathcal{F}^{-1}[(\mathcal{F}_g q_{\underline{k}_n})(X_j, \cdot)])(z) (\mathcal{F}^{-1}[(\mathcal{F}_g q_{\underline{k}_n})(X_j, \cdot)])(z)' \varpi(z) dz (\widehat{\beta}_n - \beta_n) \end{aligned}$$

and thus, following the proof of Theorem 2.1 we obtain  $A_{n2} = o_p(\sqrt{m_n})$ . Similarly as in the proof of Theorem 2.1 it can be seen that  $II_n = o_p(\sqrt{m_n})$ , which completes the proof.  $\square$

**PROOF OF PROPOSITION 6.** We decompose our test statistic as

$$\begin{aligned} nS_n &= \sum_j \int \left| \partial_z p_{\underline{m}_n}(X_j, z)' (\widehat{\alpha}_n - E[Y p_{\underline{m}_n}(X, Z)]) \right|^2 \varpi(z) dz \\ &+ 2 \sum_j \int \left( \partial_z p_{\underline{m}_n}(X_j, z)' \widehat{\alpha}_n - (\Pi_{m_n} \psi)(X_j, z) \right) \left( (\Pi_{m_n} \psi)(X_j, z) - \widehat{\psi}_{1n}(X_{1j}, z - X'_{2j} \widehat{b}_{2n}) \right) \varpi(z) dz \\ &+ \sum_j \int \left| (\Pi_{m_n} \psi)(X_j, z) - \widehat{\psi}_{1n}(X_{1j}, z - X'_{2j} \widehat{b}_{2n}) \right|^2 \varpi(z) dz \\ &= I_n + 2II_n + III_n \quad (\text{say}). \end{aligned}$$

Consider  $I_n$ . As in the proof of Proposition 5 we obtain

$$I_n = \lambda_n^{-1} \sum_{l=1}^{m_n} \tau_l \left| n^{-1/2} \sum_j \left( Y_j - \int 1\{Z_j \geq X'_1 b_1 + X'_2 b_2\} f_{B_1}(b_1) db_1 \right) p_l(X_j, Z_j) \right|^2 + o_p(1),$$

and thus, Lemma 6.2 yields  $(\sqrt{2} \varsigma_{m_n})^{-1} (I_n - \mu_{m_n}) \xrightarrow{d} \mathcal{N}(0, 1)$ . Concerning  $III_n$ , we calculate

$$\begin{aligned} III_n &\lesssim \sum_j \int \left| (\Pi_{m_n} \psi)(X_j, z) - \psi(X_j, z) \right|^2 \varpi(z) dz \\ &+ \sum_j \int \left| \psi(X_{1j}, z - X'_{2j} b_2) - \widehat{\psi}_n(X_j, z) \right|^2 \varpi(z) dz \\ &+ \sum_j \int \left| \widehat{\psi}_n(X_j, z) - \widehat{\psi}_{1n}(X_{1j}, z - X'_{2j} \widehat{b}_{2n}) \right|^2 \varpi(z) dz. \end{aligned}$$

The definition of the estimator  $\widehat{b}_2$  in (3.4) yields

$$\begin{aligned}
& \sum_j \int \left| \widehat{\psi}_n(X_j, z) - \widehat{\psi}_{1n}(X_{1j}, z - X'_{2j} \widehat{b}_{2n}) \right|^2 \varpi(z) dz \\
& \leq \sum_j \int \left| \widehat{\psi}_n(X_j, z) - \widehat{\psi}_{1n}(X_{1j}, z - X'_{2j} b_2) \right|^2 \varpi(z) dz \\
& \lesssim \sum_j \int \left| \widehat{\psi}_n(X_j, z) - \psi(X_j, z) \right|^2 \varpi(z) dz \\
& + \sum_j \int \left| \widehat{\psi}_{1n}(X_{1j}, z - X'_{2j} b_2) - \psi(X_{1j}, z - X'_{2j} b_2) \right|^2 \varpi(z) dz \\
& = o_p(\sqrt{m_n}).
\end{aligned}$$

It thus follows  $III_n = o_p(\sqrt{m_n})$ . Similarly as in the proof of Theorem 2.3 it can be shown that  $II_n = o_p(\sqrt{m_n})$ .  $\square$

## Technical Appendix.

**Lemma 6.2.** *Let Assumptions 1–3 hold true. Then*

$$(\sqrt{2}\varsigma_{m_n})^{-1} \left( \lambda_n^{-1} \sum_{l=1}^{m_n} \int \left| n^{-1/2} \sum_j \delta(V_j, t) p_l(X_j) \right|^2 \varpi(t) dt - \mu_{m_n} \right) \xrightarrow{d} \mathcal{N}(0, 1).$$

*Proof.* Let us denote the real and imaginary parts of  $\delta(V, t) p_l(X)$  by  $\delta_l^R(V, t) = \text{Re}(\delta(V, t)) p_l(X)$  and  $\delta_l^I(V, t) = \text{Im}(\delta(V, t)) p_l(X)$ , respectively. We have

$$\begin{aligned}
& \sum_{l=1}^{m_n} \int \left| (\lambda_n n)^{-1/2} \sum_j \delta(V_j, t) p_l(X_j) \right|^2 \varpi(t) dt \\
& = \sum_{l=1}^{m_n} \int \left\| (\lambda_n n)^{-1/2} \sum_j \left( \delta_l^R(V_j, t), \delta_l^I(V_j, t) \right)' \right\|^2 \varpi(t) dt \\
& = (\lambda_n n)^{-1} \sum_{l=1}^{m_n} \sum_j \int \left\| \left( \delta_l^R(V_j, t), \delta_l^I(V_j, t) \right)' \right\|^2 \varpi(t) dt \\
& + (\lambda_n n)^{-1} \sum_{l=1}^{m_n} \sum_{j \neq j'} \int \left( \delta_l^R(V_j, t) \delta_l^R(V_{j'}, t) + \delta_l^I(V_j, t) \delta_l^I(V_{j'}, t) \right) \varpi(t) dt \\
& = I_n + II_n \quad (\text{say}).
\end{aligned}$$



We observe

$$\begin{aligned}
E|I_n - \mu_{m_n}|^2 &= \text{Var}\left((\lambda_n n)^{-1} \sum_{l=1}^{m_n} \sum_j \int \left| \delta(V_j, t) p_l(X_j) \right|^2 \varpi(t) dt\right) \\
&\leq \lambda_n^{-2} n^{-1} E\left[ \int |\delta(V, t)|^4 \varpi(t) dt \left( \sum_{l=1}^{m_n} p_l^2(X) \right)^2 \right] \\
&\lesssim \sup_{x \in \mathcal{X}} \|p_{m_n}(x)\|^2 \lambda_n^{-2} n^{-1} \sum_{l=1}^{m_n} E[p_l^2(X)] \lesssim m_n^2 n^{-1} \lambda_n^{-1} = o(1)
\end{aligned}$$

using that  $\int \sup_v |\delta(v, t)|^4 \varpi(t) dt$  is bounded. Consider  $II_n$ . Let us introduce the Martingale difference array  $Q_{nj} = \sqrt{2}(\varsigma_{m_n} n)^{-1} \sum_{l=1}^{m_n} \sum_{j'=1}^{j-1} \int (\delta_l^R(V_j, t) \delta_l^R(V_{j'}, t) + \delta_l^I(V_j, t) \delta_l^I(V_{j'}, t)) \varpi(t) dt$  for  $j = 2, \dots, n$ , and zero otherwise. Further,

$$(\sqrt{2}\varsigma_{m_n})^{-1} II_n = \sqrt{2}(\varsigma_{m_n} n)^{-1} \sum_{j < j'} \sum_{l=1}^{m_n} \int (\delta_l^R(V_j, t) \delta_l^R(V_{j'}, t) + \delta_l^I(V_j, t) \delta_l^I(V_{j'}, t)) \varpi(t) dt = \sum_j Q_{nj}.$$

It remains to show that  $\sum_j Q_{nj} \xrightarrow{d} \mathcal{N}(0, 1)$ , which follows by Lemma A.3 of Breunig [2015b] by using the following computations. To show  $\sum_{j=1}^{\infty} E|Q_{nj}|^2 \leq 1$  observe that

$$\sum_{j \neq j'} \int \left( \delta_l^I(V_j, t) \delta_l^R(V_{j'}, t) - \delta_l^R(V_j, t) \delta_l^I(V_{j'}, t) \right) \varpi(t) dt = 0$$

and  $E[X_{1j} X_{1j'}] = 0$  for  $j \neq j'$ . Thus, for  $j = 2, \dots, n$  we have

$$\begin{aligned}
E|Q_{nj}|^2 &= \frac{2(j-1)}{n^2 \varsigma_{m_n}^2} E \left| \sum_{l=1}^{m_n} \int \delta_l(V_1, t) \overline{\delta_l(V_2, t)} \varpi(t) dt \right|^2 \\
&= \frac{2(j-1)}{n^2 \varsigma_{m_n}^2} \sum_{l, l'=1}^{m_n} \int \int E \left[ \delta_l(V, s) \overline{\delta_{l'}(V, t)} \right] E \left[ \overline{\delta_l(V, s)} \delta_{l'}(V, t) \right] \varpi(s) ds \varpi(t) dt \\
&= \frac{2(j-1)}{n^2 \varsigma_{m_n}^2} \sum_{l, l'=1}^{m_n} \int \int \left| E \left[ \delta_l(V, s) \overline{\delta_{l'}(V, t)} \right] \right|^2 \varpi(s) ds \varpi(t) dt \\
&= \frac{2(j-1)}{n^2}
\end{aligned}$$

by the definition of  $\varsigma_{m_n}$  and thus  $\sum_j E|Q_{nj}|^2 = 1 - 1/n$ . □

Recall  $\widehat{A}_n = n^{-1} \int \mathbf{F}_n(t)' \mathbf{F}_n(t) \varpi(t) dt$  and  $A_n = \int E[(\mathcal{F}_g p_{k_n})(X, t) (\mathcal{F}_g p_{k_n})(X, t)'] \varpi(t) dt$ .

**Lemma 6.3.** *Under the conditions of Theorem 2.1 it holds*

$$\widehat{A}_n^- = A_n^- + O_p\left(\sqrt{(\log n)k_n/n}\right).$$

*Proof.* On the set  $\Omega \equiv \left\{ \|A_n^-\| \|\widehat{A}_n - A_n\| < 1/4, \text{ rank}(A_n) = \text{rank}(\widehat{A}_n) \right\}$ , it holds  $R(\widehat{A}_n) \cap R(A_n)^\perp = \{0\}$  by Corollary 3.1 of Chen et al. [1996], where  $R$  denotes the range of a mapping. Consequently, by using properties of the Moore-Penrose pseudoinverse it holds on the set  $\Omega$ :

$$\begin{aligned} \widehat{A}_n^- - A_n^- &= -\widehat{A}_n^-(\widehat{A}_n - A_n)A_n^- + \widehat{A}_n^-(\widehat{A}_n^-)'(\widehat{A}_n - A_n)'(I_{k_n} - A_n A_n^-) \\ &\quad + (I_{k_n} - \widehat{A}_n \widehat{A}_n^-)(\widehat{A}_n - A_n)'(A_n^-)'A_n^-, \end{aligned}$$

see derivation of equation (3.19) in Theorem 3.10 on page 345 of Nashed [2014]. Applying the operator norm and using the fact that  $I_{k_n} - A_n A_n^-$  and  $I_{k_n} - \widehat{A}_n \widehat{A}_n^-$  as projections have operator norm bounded by one, we obtain

$$\begin{aligned} \|\widehat{A}_n^- - A_n^-\|_{1\Omega} &= (\|\widehat{A}_n^-\| \|\widehat{A}_n - A_n\| \|A_n^-\| + \|\widehat{A}_n^-\|^2 \|\widehat{A}_n - A_n\| + \|A_n^-\|^2 \|\widehat{A}_n - A_n\|)_{1\Omega} \\ &\leq 3 \|\widehat{A}_n - A_n\| \max \left\{ \|A_n^-\|^2, \|\widehat{A}_n^-\|^2 \right\}_{1\Omega}. \end{aligned}$$

By Theorem 3.2 of Chen et al. [1996] it holds  $\|\widehat{A}_n^-\|_{1\Omega} \leq 3\|A_n^-\| = O(1)$ . Consequently, Lemma 6.2 of Belloni et al. [2015] yields  $\|\widehat{A}_n^- - A_n^-\|_{1\Omega} = O_p(\sqrt{k_n(\log n)/n})$ . The assertion follows by  $1_\Omega = 1$  with probability approaching one.  $\square$

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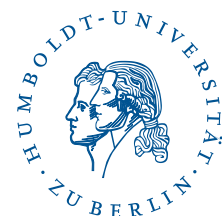
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