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Pötscher, Benedikt M. and Preinerstorfer, David

Department of Statistics, University of Vienna, ECARES, Université
libre de Bruxelles

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Valid Heteroskedasticity Robust Testing*

Benedikt M. Pötscher and David Preinerstorfer
Department of Statistics, University of Vienna
ECARES, Université libre de Bruxelles

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Abstract

Tests based on heteroskedasticity robust standard errors are an important technique in econometric practice. Choosing the right critical value, however, is not all that simple: Conventional critical values based on asymptotics often lead to severe size distortions; and so do existing adjustments including the bootstrap. To avoid these issues, we suggest to use smallest size-controlling critical values, the generic existence of which we prove in this article. Furthermore, sufficient and often also necessary conditions for their existence are given that are easy to check. Granted their existence, these critical values are the canonical choice: larger critical values result in unnecessary power loss, whereas smaller critical values lead to over-rejections under the null hypothesis, make spurious discoveries more likely, and thus are invalid. We suggest algorithms to numerically determine the proposed critical values and provide implementations in accompanying software. Finally, we numerically study the behavior of the proposed testing procedures, including their power properties.

1 Introduction

Testing hypotheses on the parameters in a regression model with potentially heteroskedastic errors is an important problem in econometrics and statistics; see MacKinnon (2013) for a recent survey. Since the classical t -statistic (F -statistic, respectively) is not pivotal (or asymptotically pivotal) in such a case in general, even under Gaussianity of the errors, so-called heteroskedasticity robust (aka heteroskedasticity consistent) modifications of these test statistics have been proposed, which are asymptotically standard normally (chi-square, respectively) distributed under the null. These modifications date back to Eicker (1963, 1967), see also Hinkley (1977), and have later been popularized in econometrics by White (1980) with great success (see MacKinnon (2013)). Unfortunately, it turned out that tests obtained from these heteroskedasticity

*Financial support of the second author by the Program of Concerted Research Actions (ARC) of the Université libre de Bruxelles is gratefully acknowledged. Address correspondence to Benedikt Pötscher, Department of Statistics, University of Vienna, A-1090 Oskar-Morgenstern Platz 1. E-Mail: benedikt.poetscher@univie.ac.at.

robust test statistics by relying on critical values derived from the respective asymptotic null distributions have a tendency to overreject the null hypothesis in finite samples (and thus are invalid), especially so if the design matrix contains leverage points; see, e.g., MacKinnon and White (1985), Davidson and MacKinnon (1985). One factor contributing to this overrejection tendency is a downward bias in the covariance matrix estimators used in these test statistics, see Chesher and Jewitt (1987). In an attempt to reduce the overrejection problem, variants of the before-mentioned heteroskedasticity robust test statistics (often denoted by HC1 through HC4, with HC0 denoting the original proposal) have been considered; see Hinkley (1977), MacKinnon and White (1985), and Cribari-Neto (2004). These variants rescale the least-squares residuals before computing the covariance matrix estimator employed in the construction of the test statistic. According to simulation studies reported in, e.g., Davidson and MacKinnon (1985) and Cribari-Neto (2004), these modifications, especially HC3 and HC4, seem to ameliorate the overrejection problem to some extent, but do not eliminate it. Further numerical results are provided in Chesher and Austin (1991), see also Chesher (1989). Numerical results in Section 9 confirm these observations. Variants of HC0-HC3, denoted by HC0R-HC3R, obtained by using restricted instead of unrestricted least-squares residuals in the computation of the covariance matrix estimators employed by the various test statistics (the restriction alluded to being the restriction defining the null hypothesis) have been introduced in Davidson and MacKinnon (1985). In their simulation experiments, this typically leads to tests that do not overreject, but that may substantially underreject; see also the simulation results in Godfrey (2006), who additionally also considers HC4R. However, as will be shown in Section 9, also these tests are in general not immune to (sometimes substantial) overrejection.

Note that all the modifications of HC0 discussed so far have the same asymptotic distribution as HC0, and thus the same critical value as for HC0 (obtained from the asymptotic null distribution) is also used for these modifications in the before mentioned literature. Sometimes small-sample adjustments to the asymptotic critical values are attempted by using the quantiles from a t_d -distribution rather than from the asymptotic normal distribution, where the degrees of freedom d are either set to $n - k$ (n and k denoting sample size and number of regressors, respectively), or are obtained through proposals set down by Satterthwaite (1946) or Bell and McCaffrey (2002); see also Imbens and Kolesár (2016). While these adjustments can lead to improvements, numerical results presented in Section 9 show that these adjustments are also not able to solve the overrejection problem in general. An alternative approach is to use bootstrap methods to compute critical values for the test statistics HC0-HC4 or HC0R-HC4R. The relevant literature is reviewed in Pötscher and Preinerstorfer (2020), and it is shown that such methods are again not immune to the overrejection problem in general.¹

A result by Bakirov and Székely (2005) needs to be mentioned here which states that – in the

¹Another possibility is to use Edgeworth expansions to find better critical values, see Rothenberg (1988) for the case of the HC0 test statistic and Davidson and MacKinnon (1985) for the HC0R test statistic. Simulation results in MacKinnon and White (1985) and Davidson and MacKinnon (1985) indicate that this does not work too well in practice. Of course, such expansions could also be worked out for the other versions of the test statistics mentioned, but this does not seem to have been pursued in the literature.

special case of testing a hypothesis on the location parameter of a heteroskedastic location model with errors that are Gaussian or scale mixtures thereof – the classical two-sided t -test (with the usual critical value) has null rejection probability not exceeding the nominal significance level under *any* form of heteroskedasticity (for a certain range of significance levels); see Ibragimov and Müller (2010) for more discussion. Ibragimov and Müller (2016), extending a result in Mickey and Brown (1966), provide a related result in the case of the comparison of two heteroskedastic populations; see also Bakirov (1998). Section 4.2 provides some more discussion. We note that all results mentioned in that section are applicable only to testing certain *scalar* linear contrasts.

Except for the Bakirov and Székely (2005) result and the variations discussed in Section 4.2, which apply only to quite special situations like, e.g., the heteroskedastic location model, none of the methods discussed so far comes with a theoretical result implying that their associated (finite sample) null rejection probabilities are guaranteed not to exceed the nominal significance level whatever the form of heteroskedasticity may be.² In fact, it transpires from the numerical results in Section 9 that for *any* of these methods instances of testing problems can be found for which the method in question overrejects substantially. Therefore, it is imperative to be able to find size-controlling critical values for the test statistics considered, i.e., critical values such that the resulting worst-case rejection probability under the null hypothesis does not exceed the nominal significance level. We shall hence pursue in this paper the construction of size-controlling critical values for the test statistics HC0-HC4, HC0R-HC4R, as well as for (two variants of) the classical (i.e., uncorrected) F -statistic (including the absolute value of the t -statistic as a special case).

In the present paper we hence consider classes of test statistics that contain the before mentioned heteroskedasticity robust test statistics as special cases and show under which conditions – and how – a critical value can be found such that the resulting test is guaranteed to have size less than or equal to α , the prescribed significance level.³ It turns out that the conditions for size controllability are broadly satisfied, in particular they are satisfied *generically* in a sense made precise further below.

We want to emphasize that the existence of size-controlling critical values for heteroskedasticity robust test statistics is *not* a trivial matter, as it has been shown in Preinerstorfer and Pötscher (2016), Section 4, that there are cases where the size of such tests is always one, *regardless* of the choice of critical value; see also the discussion in Proposition 4.7 further below. And even in cases where size control is possible by an appropriate choice of critical value, the standard critical values proposed in the literature (including the small-sample adjustments discussed

²In the special case where the number of restrictions tested equals the number of regression parameters, Davidson and Flachaire (2008) have a result which implies that certain wild bootstrap-based heteroskedasticity robust tests have size equal to the nominal significance level (and hence do not overreject) in finite samples. We note that this result in Davidson and Flachaire (2008) is not entirely correct as stated, but needs some amendments and corrections; see Pötscher and Preinerstorfer (2020).

³A less principled attempt at finding a valid test in a *given* testing problem (i.e., for given design matrix and restriction to be tested) could consist in the practitioner studying the size of a handful of tests (obtained from a few of the above mentioned test statistics in conjunction with a few of the proposed critical values) by means of an extensive Monte Carlo study and in hoping that one of the test procedures emerges from this study as valid for the particular testing problem at hand. Besides being a numerically costly procedure, it does not come with any guarantee of success.

above) are *not* guaranteed to deliver size control; in fact, they may fail to do so by a considerable margin as shown in Section 9. Our theoretical results also show the existence of a computable "threshold" C^* , say, such that any critical value C satisfying $C < C^*$ necessarily leads to a test with size 1; see Proposition 4.5. Since C^* is not difficult to compute, it can be used as a simple check to weed out unsuitable proposals for critical values.

Apart from avoiding overrejection by construction, the use of smallest size-controlling, rather than conventional, critical values offers also advantages in terms of power in instances where conventional critical values lead to underrejection (i.e., lead to a worst-case rejection probability under the null hypothesis less than the nominal significance level) as is sometimes the case; see Sections 5.1.2 and 9.2. In fact, once one has decided on a test statistic to be used for the given null hypothesis, using the smallest size-controlling critical value (provided it exists) is obviously the optimal way to proceed.

We furthermore also discuss how the critical values that lead to size control can be determined numerically and provide the R-package **hrt** (Preinerstorfer (2021)) for their computation. The usefulness of the proposed algorithms and their implementation in the R-package are illustrated numerically on some testing problems in Section 9. In particular, we compare tests obtained from various of the above mentioned test statistics when used with size-controlling critical values in terms of their power functions.

In the paper we work under a Gaussianity assumption. We stress, however, that this assumption is mainly made for convenience of presentation rather than for any other reason; as shown in Section 6.1, this assumption can be relaxed considerably.

While a trivial remark, we would like to note that the size control results given in this paper can easily be translated into results stating that the minimal coverage probability of the associated confidence set obtained by "inverting" the test is not less than the nominal confidence level.

The paper is organized as follows: After introducing notation and the test statistics in Sections 2 and 3, our size-control results are presented in Sections 4 and 5, with some further results relegated to Appendix A. Section 6 discusses ways of relaxing the underlying assumptions. Possible extensions to other classes of test statistics are discussed in Section 7, while a few comments on power are collected in Section 8. Section 9 provides the numerical results including a power study, with some details relegated to Appendix F. Section 10 concludes. Proofs and some technical results can be found in Appendices B-D. The algorithms for computing the size-controlling critical values are presented in Appendix E.

2 Framework

Consider the linear regression model

$$\mathbf{Y} = X\beta + \mathbf{U}, \tag{1}$$

where X is a (real) nonstochastic regressor (design) matrix of dimension $n \times k$ and where $\beta \in \mathbb{R}^k$ denotes the unknown regression parameter vector. We always assume $\text{rank}(X) = k$ and $1 \leq k < n$. We furthermore assume that the $n \times 1$ disturbance vector $\mathbf{U} = (\mathbf{u}_1, \dots, \mathbf{u}_n)'$ has mean zero and unknown covariance matrix $\sigma^2 \Sigma$, where Σ varies in a user-specified (nonempty) set \mathfrak{C} describing the allowed forms of heteroskedasticity, with \mathfrak{C} satisfying $\mathfrak{C} \subseteq \mathfrak{C}_{Het}$ and where $0 < \sigma^2 < \infty$ holds (σ always denoting the positive square root).⁴ The set \mathfrak{C} will be referred to as the “heteroskedasticity model”. Here

$$\mathfrak{C}_{Het} = \left\{ \text{diag}(\tau_1^2, \dots, \tau_n^2) : \tau_i^2 > 0 \text{ for all } i, \sum_{i=1}^n \tau_i^2 = 1 \right\},$$

where $\text{diag}(\tau_1^2, \dots, \tau_n^2)$ denotes the $n \times n$ matrix with diagonal elements given by τ_i^2 . That is, the errors in the regression model are uncorrelated but can be heteroskedastic. In particular, if \mathfrak{C} is chosen to be \mathfrak{C}_{Het} , one allows for heteroskedasticity of completely unknown form. The normalization condition $\sum_{i=1}^n \tau_i^2 = 1$ is included here only in order to guarantee identifiability of σ^2 and Σ , and could be replaced by any other normalization condition such as, e.g., $\max \tau_i^2 = 1$, or $\tau_1^2 = 1$, without affecting the final results (because any of these normalizations leads to the same overall set of covariance matrices $\sigma^2 \Sigma$ when σ^2 varies through the positive real line). Although a trivial observation, we stress the fact that all conceivable forms of heteroskedasticity, including parametric ones, can (possible after normalization) be cast as submodels \mathfrak{C} of \mathfrak{C}_{Het} .

Mainly for ease of exposition, we shall maintain in the sequel that the disturbance vector \mathbf{U} is normally distributed. This assumption can be substantially relaxed as discussed in Section 6.1. The linear model described in (1), together with the just made Gaussianity assumption on \mathbf{U} and with the given heteroskedasticity model \mathfrak{C} , then induces a collection of distributions on the Borel-sets of \mathbb{R}^n , the sample space of \mathbf{Y} . Denoting a Gaussian probability measure with mean $\mu \in \mathbb{R}^n$ and (possibly singular) covariance matrix A by $P_{\mu, A}$, the induced collection of distributions is then given by

$$\{P_{\mu, \sigma^2 \Sigma} : \mu \in \text{span}(X), 0 < \sigma^2 < \infty, \Sigma \in \mathfrak{C}\}. \quad (2)$$

Since every $\Sigma \in \mathfrak{C}$ is positive definite by assumption, each element of the set in the previous display is absolutely continuous with respect to (w.r.t.) Lebesgue measure on \mathbb{R}^n .

We shall consider the problem of testing a linear (better: affine) hypothesis on the parameter vector $\beta \in \mathbb{R}^k$, i.e., the problem of testing the null $R\beta = r$ against the alternative $R\beta \neq r$, where R is a $q \times k$ matrix always of rank $q \geq 1$ and $r \in \mathbb{R}^q$. Set $\mathfrak{M} = \text{span}(X)$. Define the affine space

$$\mathfrak{M}_0 = \{\mu \in \mathfrak{M} : \mu = X\beta \text{ and } R\beta = r\}$$

⁴Since we are concerned with finite-sample results only, the elements of \mathbf{Y} , X , and \mathbf{U} (and even the probability space supporting \mathbf{Y} and \mathbf{U}) may depend on sample size n , but this will not be expressed in the notation. Furthermore, the obvious dependence of \mathfrak{C} on n will also not be shown in the notation.

and let

$$\mathfrak{M}_1 = \{\mu \in \mathfrak{M} : \mu = X\beta \text{ and } R\beta \neq r\}.$$

Adopting these definitions, this testing problem can then be written more precisely as

$$H_0 : \mu \in \mathfrak{M}_0, 0 < \sigma^2 < \infty, \Sigma \in \mathfrak{C} \quad \text{vs.} \quad H_1 : \mu \in \mathfrak{M}_1, 0 < \sigma^2 < \infty, \Sigma \in \mathfrak{C}. \quad (3)$$

With \mathfrak{M}_0^{lin} we shall denote the linear space parallel to \mathfrak{M}_0 , i.e., $\mathfrak{M}_0^{lin} = \mathfrak{M}_0 - \mu_0 = \{X\beta : R\beta = 0\}$ where $\mu_0 \in \mathfrak{M}_0$. Of course, \mathfrak{M}_0^{lin} does not depend on the choice of $\mu_0 \in \mathfrak{M}_0$.

As already mentioned, the assumption of Gaussianity is made mainly for simplicity of presentation and can be relaxed substantially; see Section 6.1. The assumption of nonstochastic regressors entails little loss of generality either, which is important to emphasize: If X is random and \mathbf{U} is conditionally on X distributed as $N(0, \sigma^2 \Sigma)$, with $\sigma^2 = \sigma^2(X) > 0$ and $\Sigma = \Sigma(X) \in \mathfrak{C} \subseteq \mathfrak{C}_{Het}$, the results of the paper apply after one conditions on X (and a similar statement applies to the generalizations to non-Gaussianity discussed in Section 6.1). See Section 6.2 for more discussion. For arguments supporting conditional inference see, e.g., Robinson (1979).

We next collect some further terminology and notation used throughout the paper. A (non-randomized) *test* is the indicator function of a Borel-set W in \mathbb{R}^n , with W called the corresponding *rejection region*. The *size* of such a test (rejection region) is – as usual – defined as the supremum over all rejection probabilities under the null hypothesis H_0 given in (3), i.e.,

$$\sup_{\mu \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}} P_{\mu, \sigma^2 \Sigma}(W).$$

In slight abuse of terminology, we shall sometimes refer to this quantity as ‘the size of W over \mathfrak{C} ’ when we want to emphasize the rôle of \mathfrak{C} . Throughout the paper we let $\hat{\beta}(y) = (X'X)^{-1}X'y$, where X is the design matrix appearing in (1) and $y \in \mathbb{R}^n$. The corresponding ordinary least-squares (OLS) residual vector is denoted by $\hat{u}(y) = y - X\hat{\beta}(y)$ and its elements are denoted by $\hat{u}_t(y)$. The elements of X are denoted by x_{ti} , while x_t and $x_{\cdot i}$ denote the t -th row and i -th column of X , respectively. For \mathcal{A} an affine subspace of \mathbb{R}^n satisfying $\mathcal{A} \subseteq \text{span}(X)$ let $\tilde{\beta}_{\mathcal{A}}(y)$ denote the restricted least-squares estimator, i.e., $X\tilde{\beta}_{\mathcal{A}}(y)$ solves

$$\min_{z \in \mathcal{A}} (y - z)'(y - z).$$

Lebesgue measure on the Borel-sets of \mathbb{R}^n will be denoted by $\lambda_{\mathbb{R}^n}$, whereas Lebesgue measure on an arbitrary affine subspace \mathcal{A} of \mathbb{R}^n (but viewed as a measure on the Borel-sets of \mathbb{R}^n) will be denoted by $\lambda_{\mathcal{A}}$, with zero-dimensional Lebesgue measure being interpreted as point mass. The set of real matrices of dimension $l \times m$ is denoted by $\mathbb{R}^{l \times m}$ (all matrices in the paper will be real matrices) and Lebesgue measure on this set equipped with its Borel σ -field is denoted by $\lambda_{\mathbb{R}^{l \times m}}$. Let B' denote the transpose of a matrix $B \in \mathbb{R}^{l \times m}$ and let $\text{span}(B)$ denote the subspace in \mathbb{R}^l spanned by its columns. For a symmetric and nonnegative definite matrix B we denote the unique symmetric and nonnegative definite square root by $B^{1/2}$. For a linear subspace \mathcal{L} of

\mathbb{R}^n we let \mathcal{L}^\perp denote its orthogonal complement and we let $\Pi_{\mathcal{L}}$ denote the orthogonal projection onto \mathcal{L} . The Euclidean norm is denoted by $\|\cdot\|$, but the same symbol is also used to denote a norm of a matrix. The j -th standard basis vector in \mathbb{R}^n is written as $e_j(n)$. Furthermore, we let \mathbb{N} denote the set of all positive integers. A sum (product, respectively) over an empty index set is to be interpreted as 0 (1, respectively). Finally, for \mathcal{A} an affine subspace of \mathbb{R}^n , let $G(\mathcal{A})$ denote the group of affine transformations $y \mapsto \delta(y - a) + a^*$ where $\delta \in \mathbb{R}$, $\delta \neq 0$, and a as well as a^* are elements of \mathcal{A} ; for more information see Section 5.1 of Preinerstorfer and Pötscher (2016).

3 Test statistics

In this section we introduce the test statistics that we shall concentrate on for most of the paper. For results pertaining to other test statistics see Section 7. The test statistic we shall consider first is a standard heteroskedasticity robust test statistic frequently considered in the literature and is given by

$$T_{Het}(y) = \begin{cases} (R\hat{\beta}(y) - r)' \hat{\Omega}_{Het}^{-1}(y) (R\hat{\beta}(y) - r) & \text{if } \det \hat{\Omega}_{Het}(y) \neq 0, \\ 0 & \text{if } \det \hat{\Omega}_{Het}(y) = 0, \end{cases} \quad (4)$$

where $\hat{\Omega}_{Het} = R\hat{\Psi}_{Het}R'$ and where $\hat{\Psi}_{Het}$ is a heteroskedasticity robust estimator as considered in Eicker (1963, 1967), which later on has found its way into the econometrics literature (e.g., White (1980)). It is of the form

$$\hat{\Psi}_{Het}(y) = (X'X)^{-1}X' \text{diag}(d_1\hat{u}_1^2(y), \dots, d_n\hat{u}_n^2(y))X(X'X)^{-1},$$

where the constants $d_i > 0$ sometimes depend on the design matrix. Typical choices for d_i suggested in the literature are $d_i = 1$, $d_i = n/(n - k)$, $d_i = (1 - h_{ii})^{-1}$, or $d_i = (1 - h_{ii})^{-2}$ where h_{ii} denotes the i -th diagonal element of the projection matrix $X(X'X)^{-1}X'$, see Long and Ervin (2000) for an overview. Another suggestion is $d_i = (1 - h_{ii})^{-\delta_i}$ for $\delta_i = \min(nh_{ii}/k, 4)$, see Cribari-Neto (2004). For the last three choices of d_i just given, we use the convention that we set $d_i = 1$ in case $h_{ii} = 1$. Note that $h_{ii} = 1$ implies $\hat{u}_i(y) = 0$ for every y , and hence it is irrelevant which real value is assigned to d_i in case $h_{ii} = 1$.⁵ The five examples for the weights d_i just given correspond to what is often called HC0-HC4 weights in the literature.

In conjunction with the test statistic T_{Het} , we shall consider the following mild assumption, which is Assumption 3 in Preinerstorfer and Pötscher (2016). As discussed further below, this assumption is in a certain sense unavoidable when using T_{Het} . It furthermore also entails that our choice of assigning $T_{Het}(y)$ the value zero in case $\hat{\Omega}_{Het}(y)$ is singular has no import on the results of the paper (because of Lemma 3.1(c) below and absolute continuity of the measures $P_{\mu, \sigma^2 \Sigma}$).

⁵In fact, $h_{ii} = 1$ is equivalent to $\hat{u}_i(y) = 0$ for every y , each of which in turn is equivalent to $e_i(n) \in \text{span}(X)$.

Assumption 1. Let $1 \leq i_1 < \dots < i_s \leq n$ denote all the indices for which $e_{i_j}(n) \in \text{span}(X)$ holds where $e_j(n)$ denotes the j -th standard basis vector in \mathbb{R}^n . If no such index exists, set $s = 0$. Let $X'(\neg(i_1, \dots, i_s))$ denote the matrix which is obtained from X' by deleting all columns with indices i_j , $1 \leq i_1 < \dots < i_s \leq n$ (if $s = 0$ no column is deleted). Then $\text{rank}(R(X'X)^{-1}X'(\neg(i_1, \dots, i_s))) = q$ holds.

Observe that this assumption only depends on X and R and hence can be checked. Obviously, a simple sufficient condition for Assumption 1 to hold is that $s = 0$ (i.e., that $e_j(n) \notin \text{span}(X)$ for all j), a generically satisfied condition. Furthermore, we introduce the matrix

$$\begin{aligned} B(y) &= R(X'X)^{-1}X' \text{diag}(\hat{u}_1(y), \dots, \hat{u}_n(y)) \\ &= R(X'X)^{-1}X' \text{diag}(e'_1(n)\Pi_{\text{span}(X)^\perp}y, \dots, e'_n(n)\Pi_{\text{span}(X)^\perp}y). \end{aligned} \quad (5)$$

The facts collected in the subsequent lemma, which is taken from Pötscher and Preinerstorfer (2020) (but see also Lemma 4.1 in Preinerstorfer and Pötscher (2016)), will be used in the sequel.

Lemma 3.1. (a) $\hat{\Omega}_{Het}(y)$ is nonnegative definite for every $y \in \mathbb{R}^n$.

(b) $\hat{\Omega}_{Het}(y)$ is singular (zero, respectively) if and only if $\text{rank}(B(y)) < q$ ($B(y) = 0$, respectively).

(c) The set \mathbf{B} given by $\{y \in \mathbb{R}^n : \text{rank}(B(y)) < q\}$ (or in view of (b) equivalently given by $\{y \in \mathbb{R}^n : \det(\hat{\Omega}_{Het}(y)) = 0\}$) is either a $\lambda_{\mathbb{R}^n}$ -null set or the entire sample space \mathbb{R}^n . The latter occurs if and only if Assumption 1 is violated, in which case the test based on T_{Het} becomes trivial, as then T_{Het} is identically zero.

(d) Under Assumption 1, the set \mathbf{B} is a finite union of proper linear subspaces of \mathbb{R}^n ; in case $q = 1$, \mathbf{B} is even a proper linear subspace itself.⁶

(e) \mathbf{B} is a closed set and contains $\text{span}(X)$. Furthermore, \mathbf{B} is $G(\mathfrak{M})$ -invariant and, in particular, $\mathbf{B} + \text{span}(X) = \mathbf{B}$ holds.

In light of Part (c) of the lemma, we see that Assumption 1 is a natural and unavoidable condition if one wants to obtain a sensible test from T_{Het} .⁷ Furthermore, note that, if $\mathbf{B} = \text{span}(X)$ is true, then Assumption 1 must be satisfied (since $\text{span}(X)$ is a $\lambda_{\mathbb{R}^n}$ -null set due to the maintained assumption $k < n$). In fact, as shown in Lemma A.3 in Pötscher and Preinerstorfer (2018), the relation $\mathbf{B} = \text{span}(X)$ holds generically in various universes of design matrices. For later use we also mention that under Assumption 1 the test statistic T_{Het} is continuous at every $y \in \mathbb{R}^n \setminus \mathbf{B}$.⁸

Next, we also consider the classical (i.e., uncorrected) F-test statistic, i.e.,

$$T_{uc}(y) = \begin{cases} (R\hat{\beta}(y) - r)' \left(\hat{\sigma}^2(y) R(X'X)^{-1} R' \right)^{-1} (R\hat{\beta}(y) - r) & \text{if } y \notin \text{span}(X), \\ 0 & \text{if } y \in \text{span}(X), \end{cases} \quad (6)$$

⁶If Assumption 1 is violated, \mathbf{B} equals \mathbb{R}^n by Part (c).

⁷If this assumption is violated then T_{Het} is identically zero, an uninteresting trivial case.

⁸If Assumption 1 is violated, then T_{Het} is constant equal to zero, and hence is trivially continuous everywhere.

where $\hat{\sigma}^2(y) = \hat{u}(y)' \hat{u}(y) / (n - k) \geq 0$ (which vanishes if and only if $y \in \text{span}(X)$). Our choice to set $T_{uc}(y) = 0$ for $y \in \text{span}(X)$ again has no import on the results in the paper, since $\text{span}(X)$ is a $\lambda_{\mathbb{R}^n}$ -null set as a consequence of the maintained assumption that $k < n$ (and since the measures $P_{\mu, \sigma^2 \Sigma}$ are absolutely continuous). For reasons of comparability with (4) we have chosen not to normalize the numerator in (6) by q , the number of restrictions to be tested, as is often done in the definition of the classical F-test statistic. This also has no import on the results as the factor $1/q$ can be absorbed into the critical value. For later use we also mention that the test statistic T_{uc} is continuous at every $y \in \mathbb{R}^n \setminus \text{span}(X)$.

Remark 3.2. (i) The test statistics T_{Het} as well as T_{uc} are $G(\mathfrak{M}_0)$ -invariant as is easily seen (with the respective exceptional sets \mathbf{B} and $\text{span}(X)$ being $G(\mathfrak{M})$ -invariant).

(ii) Both statistics are actually special cases of the class of nonsphericity-corrected F-type test statistics in the sense of Section 5.4 in Preinerstorfer and Pötscher (2016) (terminology being somewhat unfortunate in case of T_{uc} as no correction for the non-sphericity is applied in this case). See Remark C.1 in Appendix C for more discussion.

We next consider two further test statistics which are versions of T_{Het} and T_{uc} with the only difference that the covariance matrix estimators used are based on restricted – instead of unrestricted – residuals. The first one of these test statistics has been suggested in the literature, e.g., in Davidson and MacKinnon (1985). We thus define

$$\tilde{T}_{Het}(y) = \begin{cases} (R\hat{\beta}(y) - r)' \tilde{\Omega}_{Het}^{-1}(y) (R\hat{\beta}(y) - r) & \text{if } \det \tilde{\Omega}_{Het}(y) \neq 0, \\ 0 & \text{if } \det \tilde{\Omega}_{Het}(y) = 0, \end{cases} \quad (7)$$

where $\tilde{\Omega}_{Het} = R\tilde{\Psi}_{Het}R'$ and where $\tilde{\Psi}_{Het}$ is given by

$$\tilde{\Psi}_{Het}(y) = (X'X)^{-1} X' \text{diag} \left(\tilde{d}_1 \tilde{u}_1^2(y), \dots, \tilde{d}_n \tilde{u}_n^2(y) \right) X(X'X)^{-1},$$

where the constants $\tilde{d}_i > 0$ sometimes depend on the design matrix and on the restriction matrix R . Here $\tilde{u}(y) = y - X\tilde{\beta}_{\mathfrak{M}_0}(y) = \Pi_{(\mathfrak{M}_0^{lin})^\perp}(y - \mu_0)$, where the last expression does not depend on the choice of $\mu_0 \in \mathfrak{M}_0$, and where $\tilde{u}_t(y)$ denotes the t -th component of $\tilde{u}(y)$. Typical choices for \tilde{d}_i are $\tilde{d}_i = 1$, $\tilde{d}_i = n/(n - (k - q))$, $\tilde{d}_i = (1 - \tilde{h}_{ii})^{-1}$, or $\tilde{d}_i = (1 - \tilde{h}_{ii})^{-2}$ where \tilde{h}_{ii} denotes the i -th diagonal element of the projection matrix $\Pi_{\mathfrak{M}_0^{lin}}$, see, e.g., Davidson and MacKinnon (1985). Another suggestion is $\tilde{d}_i = (1 - \tilde{h}_{ii})^{-\tilde{\delta}_i}$ for $\tilde{\delta}_i = \min(n\tilde{h}_{ii}/(k - q), 4)$ with the convention that $\tilde{\delta}_i = 0$ if $k = q$.⁹ For the last three choices of \tilde{d}_i just given we use the convention that we set $\tilde{d}_i = 1$ in case $\tilde{h}_{ii} = 1$. Note that $\tilde{h}_{ii} = 1$ implies $\tilde{u}_i(y) = 0$ for every y , and hence it is irrelevant which real value is assigned to \tilde{d}_i in case $\tilde{h}_{ii} = 1$.¹⁰ The five examples for the weights \tilde{d}_i just given correspond to what is often called HC0R-HC4R weights in the literature.¹¹

⁹Note that in case $k = q$ we have $\tilde{h}_{ii} = 0$, and hence $\tilde{d}_i = 1$ regardless of our convention for $\tilde{\delta}_i$.

¹⁰In fact, $\tilde{h}_{ii} = 1$ is equivalent to $\tilde{u}_i(y) = 0$ for every y , each of which in turn is equivalent to $e_i(n) \in \mathfrak{M}_0^{lin}$.

¹¹In the case $k = q$ the HC0R-HC4R weights all coincide ($\tilde{d}_i = 1$ for every i), and hence result in the same test statistic.

The subsequent assumption ensures that the set of y 's for which $\tilde{\Omega}_{Het}(y)$ is singular is a Lebesgue null set, implying that our choice of assigning $\tilde{T}_{Het}(y)$ the value zero in case $\tilde{\Omega}_{Het}(y)$ is singular has no import on the results of the paper (as the measures $P_{\mu, \sigma^2 \Sigma}$ are absolutely continuous). As discussed in the lemma further below, this assumption is in a certain sense unavoidable when using \tilde{T}_{Het} .

Assumption 2. Let $1 \leq i_1 < \dots < i_s \leq n$ denote all the indices for which $e_{i_j}(n) \in \mathfrak{M}_0^{lin}$ holds where $e_j(n)$ denotes the j -th standard basis vector in \mathbb{R}^n . If no such index exists, set $s = 0$. Let $X'(-i_1, \dots, i_s)$ denote the matrix which is obtained from X' by deleting all columns with indices i_j , $1 \leq i_1 < \dots < i_s \leq n$ (if $s = 0$ no column is deleted). Then $\text{rank}(R(X'X)^{-1}X'(-i_1, \dots, i_s)) = q$ holds.

Observe that this assumption only depends on X and R and hence can be checked. Obviously, a simple sufficient condition for Assumption 2 to hold is that $s = 0$ (i.e., that $e_j(n) \notin \mathfrak{M}_0^{lin}$ for all j), a generically satisfied condition. Furthermore, we introduce the matrix

$$\begin{aligned} \tilde{B}(y) &= R(X'X)^{-1}X' \text{diag}(\tilde{u}_1(y), \dots, \tilde{u}_n(y)) \\ &= R(X'X)^{-1}X' \text{diag}\left(e'_1(n)\Pi_{(\mathfrak{M}_0^{lin})^\perp}(y - \mu_0), \dots, e'_n(n)\Pi_{(\mathfrak{M}_0^{lin})^\perp}(y - \mu_0)\right). \end{aligned} \quad (8)$$

Note that this matrix does not depend on the choice of $\mu_0 \in \mathfrak{M}_0$. The following lemma collects some important properties of $\tilde{\Omega}_{Het}$ and \tilde{B} (defined in that lemma) and is reproduced from Pötscher and Preinerstorfer (2020) for ease of reference.

Lemma 3.3. (a) $\tilde{\Omega}_{Het}(y)$ is nonnegative definite for every $y \in \mathbb{R}^n$.

(b) $\tilde{\Omega}_{Het}(y)$ is singular (zero, respectively) if and only if $\text{rank}(\tilde{B}(y)) < q$ ($\tilde{B}(y) = 0$, respectively).

(c) The set \tilde{B} given by $\{y \in \mathbb{R}^n : \text{rank}(\tilde{B}(y)) < q\}$ (or, in view of (b), equivalently given by $\{y \in \mathbb{R}^n : \det(\tilde{\Omega}_{Het}(y)) = 0\}$) is either a $\lambda_{\mathbb{R}^n}$ -null set or the entire sample space \mathbb{R}^n . The latter occurs if and only if Assumption 2 is violated (in which case the test based on \tilde{T}_{Het} becomes trivial, as then \tilde{T}_{Het} is identically zero).

(d) Under Assumption 2, the set \tilde{B} is a finite union of proper affine subspaces; in case $q = 1$, \tilde{B} is even a proper affine subspace itself.¹² [If $r = 0$, i.e., if \mathfrak{M}_0 is linear, then the affine subspaces mentioned in the preceding sentence are in fact linear subspaces.]

(e) \tilde{B} is a closed set and contains \mathfrak{M}_0 . Also \tilde{B} is $G(\mathfrak{M}_0)$ -invariant, and in particular $\tilde{B} + \mathfrak{M}_0^{lin} = \tilde{B}$.

In light of Part (c) of the lemma, we see that Assumption 2 is a natural and unavoidable condition if one wants to obtain a sensible test from \tilde{T}_{Het} .¹³ Furthermore, note that if $\tilde{B} = \mathfrak{M}_0$ is true, then Assumption 2 must be satisfied (since \mathfrak{M}_0 is a $\lambda_{\mathbb{R}^n}$ -null set as $k - q < n$ is always the

¹²If Assumption 2 is violated, then $\tilde{B} = \mathbb{R}^n$ in view of Part (c).

¹³If this assumption is violated then \tilde{T}_{Het} is identically zero, an uninteresting trivial case.

case). For later use we also mention that under Assumption 2 the statistic \tilde{T}_{Het} is continuous at every $y \in \mathbb{R}^n \setminus \tilde{\mathbf{B}}$.¹⁴

Remark 3.4. Under the assumptions of Part (d) of Lemma 3.3 we even have that $\tilde{\mathbf{B}} - \mu_0$ is a finite union of proper linear subspaces for every $\mu_0 \in \mathfrak{M}_0$ (with the set $\tilde{\mathbf{B}} - \mu_0$ being independent of the choice of $\mu_0 \in \mathfrak{M}_0$). [This follows immediately from the result for $r = 0$ mentioned in Part (d) of the lemma upon noting that $\tilde{\mathbf{B}} - \mu_0$ coincides with the set that is constructed in the same way as $\tilde{\mathbf{B}}$ is, but from the testing problem which has $r = 0$.]

We finally consider for completeness, and in analogy with T_{uc} ,

$$\tilde{T}_{uc}(y) = \begin{cases} (R\hat{\beta}(y) - r)' \left(\tilde{\sigma}^2(y) R (X'X)^{-1} R' \right)^{-1} (R\hat{\beta}(y) - r) & \text{if } y \notin \mathfrak{M}_0, \\ 0 & \text{if } y \in \mathfrak{M}_0, \end{cases} \quad (9)$$

where $\tilde{\sigma}^2(y) = \tilde{u}(y)' \tilde{u}(y) / (n - (k - q)) \geq 0$ (which vanishes if and only if $y \in \mathfrak{M}_0$). Of course, our choice to set $\tilde{T}_{uc}(y) = 0$ for $y \in \mathfrak{M}_0$ again has no import on the results in the paper, since \mathfrak{M}_0 is a $\lambda_{\mathbb{R}^n}$ -null set. For later use we also mention that \tilde{T}_{uc} is continuous at every $y \in \mathbb{R}^n \setminus \mathfrak{M}_0$. As we shall see in Section 5.1.1, there is a close connection between \tilde{T}_{uc} and T_{uc} .

Remark 3.5. The test statistics \tilde{T}_{Het} as well as \tilde{T}_{uc} are $G(\mathfrak{M}_0)$ -invariant as is easily seen (with the respective exceptional sets $\tilde{\mathbf{B}}$ and \mathfrak{M}_0 also being $G(\mathfrak{M}_0)$ -invariant), but typically they are *not* nonsphericity-corrected F-type tests in the sense of Section 5.4 in Preinerstorfer and Pötscher (2016).

Remark 3.6. For later use we note the following: Suppose (R, r) and (\bar{R}, \bar{r}) are both of dimension $q \times (k + 1)$ and have $\text{rank}(R) = \text{rank}(\bar{R}) = q$. (i) Then (R, r) and (\bar{R}, \bar{r}) give rise to the same set \mathfrak{M}_0 , and thus to the same testing problem (3), if and only if $(AR, Ar) = (\bar{R}, \bar{r})$ holds for a nonsingular $q \times q$ matrix A . (ii) The test statistics T_{Het} , T_{uc} , \tilde{T}_{Het} , and \tilde{T}_{uc} remain the same whether they are computed using (R, r) or (\bar{R}, \bar{r}) provided $(AR, Ar) = (\bar{R}, \bar{r})$ holds for a nonsingular $q \times q$ matrix A . [To see this note that the respective exceptional sets \mathbf{B} , $\text{span}(X)$, $\tilde{\mathbf{B}}$, and \mathfrak{M}_0 are the same irrespective whether (R, r) or (\bar{R}, \bar{r}) is used, and that A cancels out in the respective quadratic forms appearing in the definition of the test statistics.]

4 Size control results for T_{Het} and T_{uc} when $\mathfrak{C} = \mathfrak{C}_{Het}$

We introduce the following notation: For a given linear subspace \mathcal{L} of \mathbb{R}^n we define the set of indices $I_0(\mathcal{L})$ via

$$I_0(\mathcal{L}) = \{i : 1 \leq i \leq n, e_i(n) \in \mathcal{L}\}.$$

We set $I_1(\mathcal{L}) = \{1, \dots, n\} \setminus I_0(\mathcal{L})$. Clearly, $\text{card}(I_0(\mathcal{L})) \leq \dim(\mathcal{L})$ holds. In particular, if $\dim(\mathcal{L}) < n$ holds (which, in particular, is so in the leading case $\mathcal{L} = \mathfrak{M}_0^{lin}$, since $\dim(\mathfrak{M}_0^{lin}) = k - q < n$), then $\text{card}(I_0(\mathcal{L})) < n$, and thus $\text{card}(I_1(\mathcal{L})) \geq 1$.

¹⁴If Assumption 2 is violated, then \tilde{T}_{Het} is constant equal to zero, and hence trivially continuous everywhere.

We have the following size control result for T_{uc} as well as for T_{Het} over the heteroskedasticity model \mathfrak{C}_{Het} (more precisely, over the null hypothesis H_0 described in (3) with $\mathfrak{C} = \mathfrak{C}_{Het}$). Note that \mathfrak{C}_{Het} is the largest possible heteroskedasticity model and reflects complete ignorance about the form of heteroskedasticity.

Theorem 4.1. (a) For every $0 < \alpha < 1$ there exists a real number $C(\alpha)$ such that

$$\sup_{\mu_0 \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(T_{uc} \geq C(\alpha)) \leq \alpha \quad (10)$$

holds, provided that

$$e_i(n) \notin \text{span}(X) \quad \text{for every } i \in I_1(\mathfrak{M}_0^{lin}). \quad (11)$$

Furthermore, under condition (11), even equality can be achieved in (10) by a proper choice of $C(\alpha)$, provided $\alpha \in (0, \alpha^*] \cap (0, 1)$ holds, where $\alpha^* = \sup_{C \in (C^*, \infty)} \sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \Sigma}(T_{uc} \geq C)$ is positive and where $C^* = \max\{T_{uc}(\mu_0 + e_i(n)) : i \in I_1(\mathfrak{M}_0^{lin})\}$ for $\mu_0 \in \mathfrak{M}_0$ (with neither α^* nor C^* depending on the choice of $\mu_0 \in \mathfrak{M}_0$).

(b) Suppose Assumption 1 is satisfied.¹⁵ Then for every $0 < \alpha < 1$ there exists a real number $C(\alpha)$ such that

$$\sup_{\mu_0 \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(T_{Het} \geq C(\alpha)) \leq \alpha \quad (12)$$

holds, provided that

$$e_i(n) \notin \mathbf{B} \quad \text{for every } i \in I_1(\mathfrak{M}_0^{lin}). \quad (13)$$

Furthermore, under condition (13), even equality can be achieved in (12) by a proper choice of $C(\alpha)$, provided $\alpha \in (0, \alpha^*] \cap (0, 1)$ holds, where now $\alpha^* = \sup_{C \in (C^*, \infty)} \sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \Sigma}(T_{Het} \geq C)$ is positive and where $C^* = \max\{T_{Het}(\mu_0 + e_i(n)) : i \in I_1(\mathfrak{M}_0^{lin})\}$ for $\mu_0 \in \mathfrak{M}_0$ (with neither α^* nor C^* depending on the choice of $\mu_0 \in \mathfrak{M}_0$).

(c) Under the assumptions of Part (a) (Part (b), respectively) implying existence of a critical value $C(\alpha)$ satisfying (10) ((12), respectively), a smallest critical value, denoted by $C_\diamond(\alpha)$, satisfying (10) ((12), respectively) exists for every $0 < \alpha < 1$. And $C_\diamond(\alpha)$ corresponding to Part (a) (Part (b), respectively) is also the smallest among the critical values leading to equality in (10) ((12), respectively) whenever such a critical value exists. [Although $C_\diamond(\alpha)$ corresponding to Part (a) and (b), respectively, will typically be different, we use the same symbol.]

We see from the theorem that the condition for size control of T_{Het} (T_{uc} , respectively) over \mathfrak{C}_{Het} , i.e., condition (13) ((11), respectively), only depends on X and R ; in particular, in case of T_{Het} , it does not depend on how the weights d_i figuring in the definition of T_{Het} have been chosen (note that the set \mathbf{B} only depends on X and R). Moreover, the sufficient conditions for size control are generically satisfied in the universe of all $n \times k$ design matrices X (of rank k),

¹⁵Condition (13) clearly implies that the set \mathbf{B} is a proper subset of \mathbb{R}^n (as $\text{card}(I_1(\mathfrak{M}_0^{lin})) \geq 1$) and thus implies Assumption 1. Hence, we could have dropped this assumption from the formulation of the theorem. For clarity of presentation we have, however, chosen to explicitly mention Assumption 1. A similar remark applies to some of the other results given below and will not be repeated.

see Example 4.1 and the attending discussion further below. Furthermore, it is plain that the size-controlling critical values $C(\alpha)$ in Theorem 4.1 will depend on the choice of test statistic as well as on the testing problem at hand. More concretely, the size-controlling critical values in Part (b) of the theorem thus depend only on X , R , and r , as well as on the choice of weights d_i , whereas in Part (a) the dependence is only on X , R , and r . We do not show these dependencies in the notation. In fact, as discussed in Remark 4.2 below, it turns out that the size-controlling critical values in both cases actually do *not* depend on the value of r at all (provided the weights d_i are not allowed to depend on r in case of T_{Het}). Similarly, it is easy to see that C^* and α^* in Theorem 4.1 do not depend on r (under the same provision as before in case of T_{Het}).

Another observation is that any critical value delivering size control over \mathfrak{C}_{Het} also delivers size control over *any* other heteroskedasticity model \mathfrak{C} since $\mathfrak{C} \subseteq \mathfrak{C}_{Het}$. Of course, for such a \mathfrak{C} even smaller critical values (than needed for \mathfrak{C}_{Het}) may already suffice for size control. Also note that sufficient conditions implying size control over \mathfrak{C}_{Het} may be more restrictive than sufficient conditions implying only size control over a smaller heteroskedasticity model \mathfrak{C} . For size control results tailored to such smaller models \mathfrak{C} see Appendix A.

In light of the results of Chesher and Jewitt (1987) and Chesher (1989), it is useful to interpret (11) and (13) (which generically coincides with (11); see Lemma A.3 in Pötscher and Preinerstorfer (2018)) in terms of leverage points. Note that $e_i(n) \in \text{span}(X)$ is equivalent to $h_{ii} = 1$, which corresponds to the i -th observation being an “extreme leverage point”. Hence, the condition for a size-controlling critical value to exist in Part (a) of Theorem 4.1 requires that none of the indices in $\mathcal{I}_1(\mathfrak{M}_0^{lin})$ corresponds to an extreme leverage point. It is interesting to observe that all indices in $\mathcal{I}_0(\mathfrak{M}_0^{lin})$ (note that this set may be empty) correspond to extreme leverage points themselves. Hence, for the condition in (11) not to be satisfied, not only must extreme leverage points exist, but the lever needs to be of a particular type depending on the hypothesis (R, r) .

Remark 4.2. (*Independence of the value of r and implications for confidence sets*) (i) As already noted before, the sufficient conditions for size control in both parts of Theorem 4.1 only depend on X and R . In particular, they do not depend on the value of r .

(ii) The size of the test based on T_{uc} (T_{Het} , respectively) in Theorem 4.1 as well as the size-controlling critical values $C(\alpha)$ (for both test statistics) do also not depend on the value of r (provided the weights d_i are not allowed to depend on r in case of T_{Het}). This follows from Lemma 5.15 in Pötscher and Preinerstorfer (2018) combined with Remark C.1 in Appendix C.¹⁶ This observation is of some importance, as it allows one easily to obtain confidence sets for $R\beta$ by “inverting” the test without the need of recomputing the critical value for every value of r .

Remark 4.3. (*Some equivalencies*) If the respective smallest size-controlling critical values are used (provided they exist), the tests obtained from T_{Het} with the HC0 and the HC1 weights, respectively, are identical, as these two test statistics differ only by a multiplicative constant.

¹⁶For this argument we impose Assumption 1 in case of T_{Het} , the case where this assumption is violated being trivial.

The same reasoning applies to the test statistics based on the HC0-HC4 weights, respectively, in case h_{ii} does not depend on i .

Remark 4.4. (*Positivity of size-controlling critical values*) For every $0 < \alpha < 1$ any $C(\alpha)$ satisfying (10) or (12) is necessarily positive. To see this observe that $\{T_{uc} \geq C\} = \{T_{Het} \geq C\} = \mathbb{R}^n$ for $C \leq 0$, since both test statistics are nonnegative everywhere.

The next proposition complements Theorem 4.1 and provides a useful lower bound for the size-controlling critical values (other than the trivial bound given in the preceding remark).

Proposition 4.5. ^{17,18} (a) Suppose that (11) is satisfied. Then any $C(\alpha)$ satisfying (10) necessarily has to satisfy $C(\alpha) \geq C^*$, where C^* is as in Part (a) of Theorem 4.1. In fact, for any $C < C^*$ we have $\sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(T_{uc} \geq C) = 1$ for every $\mu_0 \in \mathfrak{M}_0$ and every $\sigma^2 \in (0, \infty)$.

(b) Suppose that Assumption 1 and (13) are satisfied. Then any $C(\alpha)$ satisfying (12) necessarily has to satisfy $C(\alpha) \geq C^*$, where C^* is as in Part (b) of Theorem 4.1. In fact, for any $C < C^*$ we have $\sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(T_{Het} \geq C) = 1$ for every $\mu_0 \in \mathfrak{M}_0$ and every $\sigma^2 \in (0, \infty)$.

The preceding observation is useful in two ways: First, critical values suggested in the literature (such as, e.g., the $(1 - \alpha)$ -quantile of a chi-square distribution with q degrees of freedom or critical values obtained from a degree of freedom adjustment) can immediately be dismissed if they turn out to be less than C^* , as they then certainly will not guarantee size control. We use this line of reasoning in the numerical results in Section 9. Second, if the *observed* value of the test statistic T_{Het} (T_{uc} , respectively) is less than C^* , the decision not to reject the null hypothesis can be taken without further need to compute size-controlling critical values. Note that C^* as given in Theorem 4.1 is quite easy to compute.

Remark 4.6. Suppose the assumptions of Part (a) (Part (b), respectively) of Theorem 4.1 are satisfied. Then we know from that theorem that the size (over \mathfrak{C}_{Het}) of $\{T_{uc} \geq C_\diamond(\alpha)\}$ ($\{T_{Het} \geq C_\diamond(\alpha)\}$, respectively) equals α provided $\alpha \in (0, \alpha^*] \cap (0, 1)$. If now $\alpha^* < \alpha < 1$, then the size (over \mathfrak{C}_{Het}) of $\{T_{uc} \geq C_\diamond(\alpha)\}$ ($\{T_{Het} \geq C_\diamond(\alpha)\}$, respectively) equals α^* (where the $C_\diamond(\alpha)$'s pertaining to Parts (a) and (b) may be different). This follows from $C_\diamond(\alpha) \geq C^*$ (see Proposition 4.5 above) and Remark 5.13(i) in Pötscher and Preinerstorfer (2018).¹⁹ This argument actually also delivers that $C_\diamond(\alpha) = C^*$ must hold in case $\alpha^* < \alpha < 1$.

¹⁷It is not difficult to show in the context of Parts (a) and (b) of the proposition that any critical value $C > C^*$ actually leads to size less than 1. This follows from a reasoning similar as in Remark 5.4 of Pötscher and Preinerstorfer (2018).

¹⁸If (13) in Part (b) of the proposition does not hold, the conclusion of Part (b) can be shown to continue to hold with C^* as defined in Theorem 4.1(b), and also with C^* as defined in Lemma 5.11 of Pötscher and Preinerstorfer (2018) (note that under the assumptions of Part (b) of the proposition both definitions of C^* actually coincide as shown in the proof of Theorem 4.1). [Recall that under violation of (13) size-controlling critical values may or may not exist.] If Assumption 1 is not satisfied, then $T_{Het} \equiv 0$, and the conclusion of Part (b) holds trivially (as $C^* = 0$ with both definitions). If (11) in Part (a) of the proposition is not satisfied, then no size-controlling critical value exists by Proposition 4.7; hence, the conclusion of Part (a) holds trivially, again regardless of which of the two definitions of C^* is adopted.

¹⁹The assumptions for Part A of Proposition 5.12 in Pötscher and Preinerstorfer (2018) required in Remark 5.13 of that paper are satisfied under the assumptions of Theorem 4.1 as shown in the proof of Theorem A.1 in Appendix C. In this proof also the condition $\lambda_{\mathbb{R}^n}(T_{uc} = C^*) = 0$ ($\lambda_{\mathbb{R}^n}(T_{Het} = C^*) = 0$, respectively) required in Remark 5.13 of Pötscher and Preinerstorfer (2018) is verified.

We next discuss to what extent the sufficient conditions for size control in Theorem 4.1 are also necessary.

Proposition 4.7. (a) *If (11) is violated, then $\sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(T_{uc} \geq C) = 1$ for every choice of critical value C , every $\mu_0 \in \mathfrak{M}_0$, and every $\sigma^2 \in (0, \infty)$ (implying that size equals 1 for every C). As a consequence, the sufficient condition for size control (11) in Part (a) of Theorem 4.1 is also necessary.*

(b) *Suppose Assumption 1 is satisfied.²⁰ If (11) is violated, then $\sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(T_{Het} \geq C) = 1$ for every choice of critical value C , every $\mu_0 \in \mathfrak{M}_0$, and every $\sigma^2 \in (0, \infty)$ (implying that size equals 1 for every C). [In case X and R are such that $\mathbf{B} = \text{span}(X)$, conditions (11) and (13) coincide; hence the sufficient condition for size control (13) in Part (b) of Theorem 4.1 is then also necessary in this case.]*

Remark 4.8. Suppose Assumption 1 is satisfied. In case $\mathbf{B} \neq \text{span}(X)$ and (11) hold, but (13) is violated, neither Part (b) of Theorem 4.1 nor Part (b) of Proposition 4.7 apply. We note that there are instances of this situation (see Example 4.5) for which it can be shown by other methods that T_{Het} is size controllable despite failure of (13);²¹ as a consequence, (13) is not necessary for (12) in general. We conjecture that there are other instances of the situation described here where size control is not possible, but we have not investigated this in any detail.²² [What can be said in general in this situation is that the size of the rejection region $\{T_{Het} \geq C\}$ over \mathfrak{C}_{Het} is certainly equal to 1 for every $C < \max \{T_{Het}(\mu_0 + e_i(n)) : e_i(n) \notin \mathbf{B}\}$, where we use the convention that this maximum is $-\infty$ in case the set over which the maximum is taken is empty. This follows from Lemma 4.1 in Pötscher and Preinerstorfer (2019) with \mathbb{K} equal to the collection $\{\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n) : e_i(n) \notin \mathbf{B}\}$.]

Remark 4.9. Let T stand for either T_{Het} or T_{uc} in the following and suppose that Assumption 1 is satisfied in case of $T = T_{Het}$: By Remark C.1 in Appendix C and Lemma 5.16 in Pötscher and Preinerstorfer (2018) the rejection regions $\{y : T(y) \geq C\}$ and $\{y : T(y) > C\}$ differ only by a $\lambda_{\mathbb{R}^n}$ -null set. Since the measures $P_{\mu, \sigma^2 \Sigma}$ are absolutely continuous w.r.t. $\lambda_{\mathbb{R}^n}$ when Σ is nonsingular, $P_{\mu, \sigma^2 \Sigma}(T \geq C) = P_{\mu, \sigma^2 \Sigma}(T > C)$ then follows, and hence the results in this section given for rejection probabilities $P_{\mu, \sigma^2 \Sigma}(T \geq C)$ apply to rejection probabilities $P_{\mu, \sigma^2 \Sigma}(T > C)$ equally well. A similar remark applies to the results in Appendix A.1.

4.1 Some examples

We illustrate Theorem 4.1 and Proposition 4.7 with a few examples.

Example 4.1. (i) Suppose the design matrix satisfies $e_i(n) \notin \text{span}(X)$ for every $1 \leq i \leq n$ (which will typically be the case). Then obviously the sufficient condition (11) is satisfied (in

²⁰If this assumption is violated then T_{Het} is identically zero, an uninteresting trivial case.

²¹In this example actually $e_i(n) \in \mathbf{B}$ holds for all $i = 1, \dots, n$.

²²In case $q = 1$, Theorem A.4 in Pötscher and Preinerstorfer (2019) could potentially be used here to derive conditions when size control is not possible.

fact, for every choice of \mathfrak{M}_0 , i.e., for every choice of restriction to be tested). And the sufficient condition (13) is also satisfied provided $\mathbf{B} = \text{span}(X)$.

(ii) Suppose the design matrix X and the restriction R are such that $e_i(n) \notin \mathbf{B}$ for *every* $1 \leq i \leq n$. Then the sufficient condition (13) is clearly satisfied.

This example shows, in particular, that the sufficient conditions for size control are generically satisfied in the universe of all $n \times k$ design matrices X (of rank k). Given the example, this is obvious for T_{uc} ; and it follows for T_{Het} by additionally noting that, for every given choice of restriction to be tested, the relation $\mathbf{B} = \text{span}(X)$ holds generically in the universe of all $n \times k$ design matrices X (of rank k); see Lemma A.3 in Pötscher and Preinerstorfer (2018). The next example discusses the case where a standard basis vector is among the regressors.

Example 4.2. Suppose that $e_1(n)$ is the first column of X and that $e_i(n) \notin \text{span}(X)$ for every $2 \leq i \leq n$. Suppose further that R is of the form $R = (0, \tilde{R})$, where \tilde{R} has dimension $q \times (k-1)$. That is, the restriction to be tested does not involve the coefficient of the first regressor. Then it is easy to see that (11) is satisfied and size control for T_{uc} is thus possible. If also $\mathbf{B} = \text{span}(X)$ holds, then the same is true for (13) and T_{Het} . [In case R is not as above, but has a nonzero first coordinate, then it is easy to see that $1 \in I_1(\mathfrak{M}_0^{lin})$, and hence (11) is violated. It follows from Proposition 4.7 that the rejection region $\{T_{uc} \geq C\}$ indeed has size 1 for every choice of critical value C when \mathfrak{C}_{Het} is the heteroskedasticity model; and the same is true for T_{Het} , provided Assumption 1 is satisfied.²³]

We continue with a few more examples where X has a particular structure.

Example 4.3. (*Heteroskedastic location model*) Suppose $k = 1$, $x_{t1} = 1$ for all t , $q = 1$, $R = 1$, and $r \in \mathbb{R}$. The heteroskedasticity model is given by \mathfrak{C}_{Het} . Then the conditions for size control in both parts of Theorem 4.1 are satisfied (since it is easy to see that \mathbf{B} coincides with $\text{span}(X)$ and that Assumption 1 is satisfied). Note also that in this example T_{Het} and T_{uc} actually coincide in case $d_i = n/(n-1)$ for all i , i.e., if the HC1, HC2, or HC4 weights are used, and differ only by a multiplicative constant if the HC0 or HC3 weights are employed; in particular, all these test statistics give rise to one and the same test if the respective smallest size-controlling critical values are used (cf. Remark 4.3).²⁴ Furthermore, note that the here observed size controllability is in line with results in Bakirov and Székely (2005) stating that, for a certain range of significance levels α , the usual critical values obtained from an $F_{1,n-1}$ -distribution actually can be used as size-controlling critical values $C(\alpha)$ for the test statistic T_{uc} (in fact, these are then the smallest size-controlling critical values $C_\diamond(\alpha)$).

The subsequent example is closely related to the Behrens-Fisher problem, see Remark A.4 in Appendix A.1.

²³If Assumption 1 is violated then T_{Het} is identically zero, an uninteresting trivial case.

²⁴In fact, more is true in the location model: The test statistics \tilde{T}_{Het} using the HC0R-HC4R weights all coincide (cf. Footnote 11), and they also coincide with \tilde{T}_{uc} . Perusing the connection between \tilde{T}_{uc} and T_{uc} established in Section 5.1.1, we can then even conclude that all the test statistics T_{uc} , T_{Het} with HC0-HC4 weights, \tilde{T}_{uc} , and \tilde{T}_{Het} with HC0R-HC4R weights give rise to (essentially) the same test, provided the respective smallest size-controlling critical values are used.

Example 4.4. (*Comparing the means of two heteroskedastic groups*) Consider the problem of testing the equality of the means of two independent normal populations where the variances of each item may be different, even within a group. In our framework this corresponds to the case $k = 2$, $x_{t1} = 1$ for $1 \leq t \leq n_1$, $x_{t1} = 0$ for $n_1 < t \leq n_1 + n_2 = n$, $x_{t2} = 1 - x_{t1}$, and $R = (1, -1)$ with $r = 0$. The heteroskedasticity model is then again \mathfrak{C}_{Het} . We first assume that $n_i \geq 2$ holds for $i = 1, 2$. Note that in the present context T_{uc} is nothing else than the square of the two-sample t-statistic that uses a pooled variance estimator, and that T_{Het} is the square of the two-sample t-statistic that uses appropriate variance estimators from each group (the particular form of the variance estimator being determined by the choice of d_i). Now, $e_i(n) \notin \text{span}(X)$ for every $1 \leq i \leq n$ holds, and hence T_{uc} is size controllable (cf. Example 4.1(i)). This is in line with results in Bakirov (1998), cf. also Section 4.2. Furthermore, it is obvious that Assumption 1 is satisfied (as $s = 0$) and a simple calculation shows that $B(y) = \hat{u}(y)'A$, where A is a diagonal matrix with $a_{ii} = n_1^{-1}$ for $1 \leq i \leq n_1$ and $a_{ii} = -n_2^{-1}$ else. This shows that the set \mathbf{B} coincides with $\text{span}(X)$. Consequently, also T_{Het} is size controllable (again cf. Example 4.1(i)). We also note here that the observed size controllability of T_{Het} is in line with results in Ibragimov and Müller (2016) stating that for a certain range of significance levels α and group sizes n_i the usual critical values obtained from an $F_{1, \min(n_1, n_2) - 1}$ -distribution actually can be used as size-controlling critical values $C(\alpha)$ for the test statistic T_{Het} in case d_i is set equal to $(1 - h_{ii})^{-1}$. In fact, they are then the smallest size-controlling critical values, cf. the discussion preceding Theorem 1 in Ibragimov and Müller (2016). In the rather uninteresting case $n_1 = 1$ and $n_2 \geq 2$, it is easy to see that Assumption 1 is satisfied and that the size of both tests equals 1 for all choices of critical values in view of Proposition 4.7, since $e_1(n) \in \text{span}(X)$ and $1 \in I_1(\mathfrak{M}_0^{lin}) = \{1, \dots, n\}$. The same is true if $n_1 \geq 2$ and $n_2 = 1$. [The remaining and uninteresting case $n_1 = n_2 = 1$ falls outside of our framework since we always require $n > k$.]

The next example is an extension of the previous problem to the case of more than two groups. An interesting phenomenon occurs here: The sufficient conditions for size control of T_{Het} given in Theorem 4.1 are *violated*, but size controllability can *nevertheless* be established by additional arguments. Hence, this example provides an instance where the conditions in Part (b) of Theorem 4.1 are not necessary.

Example 4.5. (*Comparing the means of k heteroskedastic groups*) We are given k integers $n_j \geq 1$ with $\sum_{j=1}^k n_j = n$ describing group sizes where $k \geq 3$ holds. The regressors x_{ti} for $1 \leq i \leq k$ indicate group membership, i.e., they satisfy $x_{ti} = 1$ for $\sum_{j=1}^{i-1} n_j < t \leq \sum_{j=1}^i n_j$ and $x_{ti} = 0$ otherwise. The heteroskedasticity model is given by \mathfrak{C}_{Het} . We are interested in testing $\beta_1 = \dots = \beta_k$. We thus may choose the $(k-1) \times k$ restriction matrix R with j -th row $(1, 0, \dots, 0, -1, \dots, 0)$ where the entry -1 is at position $j+1$. Of course, $q = k-1$ and $r = 0$ hold. We first consider the case where $n_j \geq 2$ for all j . Then clearly $k < n$ is satisfied. With regard to T_{uc} we see immediately that $e_i(n) \notin \text{span}(X)$ for every $1 \leq i \leq n$ follows (since $n_j \geq 2$ for all j) and thus the sufficient condition (11) for size control of T_{uc} is satisfied. Turning to T_{Het} , it is easy to see that Assumption 1 is satisfied (since $s = 0$ in view of $n_j \geq 2$). Furthermore, the j -th

row of $R(X'X)^{-1}X'$ is seen to be of the form

$$(n_1^{-1}, \dots, n_1^{-1}, 0, \dots, 0, -n_{j+1}^{-1}, \dots, -n_{j+1}^{-1}, 0, \dots, 0),$$

from which it follows that

$$R(X'X)^{-1}X' \text{diag}(d_1 \hat{u}_1^2(y), \dots, d_n \hat{u}_n^2(y))X(X'X)^{-1}R = S_1 \iota' + \text{diag}(S_2, \dots, S_k), \quad (14)$$

where ι is the $(k-1)$ -dimensional vector with entries all equal to 1 and where $S_j = n_j^{-2} \sum_t d_t \hat{u}_t^2(y) = n_j^{-2} \sum_t d_t (y_t - \bar{y}_{(j)})^2$ with the summation index t running over all elements in the j -th group, and where $\bar{y}_{(j)}$ is the mean in group j . From (14) it is not difficult to verify that the set \mathbf{B} is given by

$$\mathbf{B} = \bigcup_{i,j=1, i \neq j}^k \{y \in \mathbb{R}^n : S_i(y) = S_j(y) = 0\} = \bigcup_{i,j=1, i \neq j}^k \text{span}(x_{\cdot i}, x_{\cdot j}, \{e_l(n) : x_{li} = x_{lj} = 0\}).$$

Note that \mathbf{B} is not a linear space and is strictly larger than $\text{span}(X)$. The set \mathfrak{M}_0^{lin} is given by the span of the vector $e = (1, 1, \dots, 1)'$. Hence, $I_1(\mathfrak{M}_0^{lin}) = \{1, \dots, n\}$. Since $e_i(n) \in \mathbf{B}$ holds for every i , we conclude that the sufficient condition (13) for size control of T_{Het} is *not* satisfied and hence Part (b) of Theorem 4.1 does *not* apply. However, it can be shown by additional arguments, see Proposition C.3 in Appendix C, that T_{Het} is nevertheless size controllable, i.e., that (12) holds. Next, in the case where $n_j = 1$ for some j , but not for all j , Proposition 4.7 shows that the size of the test based on T_{uc} equals 1 for all choices of critical values, since then for some i the standard basis vector $e_i(n)$ is one of the regressors and thus we have $e_i(n) \in \text{span}(X)$ and $i \in I_1(\mathfrak{M}_0^{lin}) = \{1, \dots, n\}$. For T_{Het} the same is true if $n_j = 1$ holds for exactly one j (because of Part (b) of Proposition 4.7 and since then Assumption 1 is satisfied as is easily seen); in case $n_j = 1$ is true for (at least) two, but not all, values of j , T_{Het} is identically zero (as then Assumption 1 is violated), and thus is size-controllable in a trivial way. [The remaining and uninteresting case $n_j = 1$ for all j falls outside of our framework since we always require $n > k$.]

We close this section by one more example. Again, the sufficient conditions in Part (b) of Theorem 4.1 fail to hold, but additional arguments based on Example 4.3 establish size controllability of the test based on T_{Het} .

Example 4.6. Consider again the situation of Example 4.4, except that now $R = I_2$, the 2×2 identity matrix, (and again $r = 0$). Consider first the case where $n_i \geq 2$ for $i = 1, 2$. Condition (11) is obviously satisfied (since $n_i \geq 2$ for $i = 1, 2$) and hence T_{uc} is size controllable. We next turn to T_{Het} . Since $\mathfrak{M}_0^{lin} = \{0\}$ we have $I_1(\mathfrak{M}_0^{lin}) = \{1, \dots, n\}$. Furthermore, simple computations show that Assumption 1 is satisfied and that

$$\mathbf{B} = \text{span}(x_{\cdot 1}, \{e_i(n) : i > n_1\}) \cup \text{span}(x_{\cdot 2}, \{e_i(n) : i \leq n_1\}).$$

Obviously, the sufficient condition (13) for size control of T_{Het} is violated. Nevertheless, T_{Het} is size controllable by the following argument: Simple computations show that $T_{Het}(y) = T_1(y) + T_2(y)$ for $y \notin B$, where $T_1(y) = n_1^2 \hat{\beta}_1^2(y) / \sum_{t=1}^{n_1} d_t \hat{u}_t^2(y)$ and $T_2(y) = n_2^2 \hat{\beta}_2^2(y) / \sum_{t=n_1+1}^n d_t \hat{u}_t^2(y)$. [If the denominator in the formula for $T_i(y)$ is zero for some $y \in \mathbb{R}^n$, we define $T_i(y)$ as zero.] Since B is a $\lambda_{\mathbb{R}^n}$ -null set, $P_{0,\sigma^2\Sigma}(T_{Het} \geq C) \leq P_{0,\sigma^2\Sigma}(T_1 \geq C/2) + P_{0,\sigma^2\Sigma}(T_2 \geq C/2)$ for $C > 0$. Now, it is easy to see that $P_{0,\sigma^2\Sigma}(T_i \geq C/2)$ for $i = 1, 2$ coincides with the null rejection probability of a test for the mean in a heteroskedastic location model (based on a test statistic of the form (4)). However, as shown in Example 4.3, such a test is size controllable. [In the case $n_1 = 1$ and $n_2 \geq 2$ (or vice versa) condition (11) is violated and the rejection region $\{T_{uc} \geq C\}$ has size 1 for every C ; furthermore, Assumption 1 is violated, and hence T_{Het} is identically zero. The case $n_1 = n_2 = 1$ falls outside of our framework as then $k = n$.]

In Appendix C we discuss yet another example where the sufficient condition of Part (b) of Theorem 4.1 fails, but size-controllability can nevertheless be established.

4.2 Some variations on Bakirov and Székely (2005)

(i) As noted in Ibragimov and Müller (2010), testing a hypothesis regarding a *scalar* linear contrast in a heteroskedastic (Gaussian) linear regression model more general than a location model can often be converted to a testing problem in a heteroskedastic (Gaussian) location model by suitably dividing the data into subgroups and by considering groupwise least-squares estimators, thus making it amenable to the Bakirov and Székely (2005) result mentioned in Section 1. However, this introduces additional questions, e.g., as how to divide up the data. In any case, this approach is limited to testing hypotheses on *scalar* linear contrasts. It also requires that the linear contrast subject to test is *estimable* in each subgroup.

(ii) In case the linear contrast subject to test is *not* estimable in each subgroup, but can be written as the difference of two linear contrasts where the first contrast is estimable in the first G_1 groups whereas the second contrast is estimable in the last G_2 groups (where we consider a total of $G_1 + G_2$ groups), Ibragimov and Müller (2016) point out that the problem can be converted into the problem of comparing two heteroskedastic (Gaussian) populations. Now, for such a two-sample comparison problem Bakirov (1998) shows for a certain two-sample t -statistic (the square of which is T_{uc} , cf. Example 4.4 above) how – in the presence of heteroskedasticity – size-controlling critical values can be constructed by appropriately transforming quantiles of a t -distribution; this result imposes conditions which entail that the nominal significance α level must be quite small (requiring α not to exceed 0.01 for many group sizes, and often to be considerably smaller). This somewhat limits the applicability of Bakirov’s result. Thus Ibragimov and Müller (2016) go on to consider another two-sample t -statistic (the square of which is T_{Het} with $d_i = (1 - h_{ii})^{-1}$, cf. Example 4.4 above) and – extending a result in Mickey and Brown (1966) – provide a Bakirov and Székely (2005)-type result, i.e., they show that the $(1 - \alpha/2)$ -quantile of a t -distribution with degrees of freedom equal to the smaller of the two sample sizes

minus 1 provides the smallest size-controlling critical value even under heteroskedasticity.²⁵ This result holds under certain conditions on the sample sizes and only for small α , but, e.g., allows for the choice $\alpha = 0.05$. [We note here that the description of Bakirov (1998)’s result in Ibragimov and Müller (2016) is inaccurate in that a certain transformation of the critical value is being ignored.]

(iii) In the problem of comparing two heteroskedastic (Gaussian) populations based on samples of equal size (“balanced design”) one can – instead of using the two-sample t -test statistics considered in Bakirov (1998) and Ibragimov and Müller (2016) – employ the Bartlett test statistic, which simply is the usual t -test statistic computed from the differences between the observations in the two samples.²⁶ An advantage of this approach is that the original Bakirov and Székely (2005) result is directly applicable, and there is no need to resort to the results described in (ii).

(iv) Another quite special case that can be brought under the realm of the Bakirov and Székely (2005) result is a heteroskedastic (Gaussian) regression model with only one regressor that never takes the value zero. Dividing the t -th equation in the regression model by x_t , converts this into a heteroskedastic location problem.

(v) The results in (i)-(iv) immediately also apply if the errors in the regression are distributed as scale mixtures of Gaussians (cf. also Section 6.1).

5 Size control results for \tilde{T}_{Het} and \tilde{T}_{uc} when $\mathfrak{C} = \mathfrak{C}_{Het}$

Here we first discuss size control results for \tilde{T}_{uc} as well as for \tilde{T}_{Het} over the heteroskedasticity model \mathfrak{C}_{Het} (more precisely, over the null hypothesis H_0 described in (3) with $\mathfrak{C} = \mathfrak{C}_{Het}$). In Subsection 5.1 we then discuss some peculiar properties of the test statistics \tilde{T}_{uc} and \tilde{T}_{Het} .

We note that the first statement in Part (a) of the subsequent theorem is actually trivial, since \tilde{T}_{uc} is bounded as shown in the next subsection (which also provides a discussion when non-trivial size-controlling critical values exist).

Theorem 5.1. (a) *For every $0 < \alpha < 1$ there exists a real number $C(\alpha)$ such that*

$$\sup_{\mu_0 \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(\tilde{T}_{uc} \geq C(\alpha)) \leq \alpha \quad (15)$$

holds. Furthermore, even equality can be achieved in (15) by a proper choice of $C(\alpha)$, provided $\alpha \in (0, \alpha^] \cap (0, 1)$ holds, where $\alpha^* = \sup_{C \in (C^*, \infty)} \sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \Sigma}(\tilde{T}_{uc} \geq C)$ and where $C^* = \max\{\tilde{T}_{uc}(\mu_0 + e_i(n)) : i \in I_1(\mathfrak{M}_0^{lin})\}$ for $\mu_0 \in \mathfrak{M}_0$ (with neither α^* nor C^* depending on the choice of $\mu_0 \in \mathfrak{M}_0$).*

²⁵In the balanced case (i.e., if the two samples have the same cardinality) the test statistic considered in Bakirov (1998) actually coincides with the test statistic in Ibragimov and Müller (2016).

²⁶Certainly, there is some arbitrariness in how the observations are being “paired”.

(b) Suppose Assumption 2 is satisfied.²⁷ Suppose further that \tilde{T}_{Het} is not constant on $\mathbb{R}^n \setminus \tilde{\mathbf{B}}$.²⁸ Then for every $0 < \alpha < 1$ there exists a real number $C(\alpha)$ such that

$$\sup_{\mu_0 \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(\tilde{T}_{Het} \geq C(\alpha)) \leq \alpha \quad (16)$$

holds, provided that for some $\mu_0 \in \mathfrak{M}_0$ (and hence for all $\mu_0 \in \mathfrak{M}_0$)

$$\mu_0 + e_i(n) \notin \tilde{\mathbf{B}} \quad \text{for every } i \in I_1(\mathfrak{M}_0^{lin}). \quad (17)$$

Furthermore, under condition (17), even equality can be achieved in (16) by a proper choice of $C(\alpha)$, provided $\alpha \in (0, \alpha^*] \cap (0, 1)$ holds, where now $\alpha^* = \sup_{C \in (C^*, \infty)} \sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \Sigma}(\tilde{T}_{Het} \geq C)$ and where $C^* = \max\{\tilde{T}_{Het}(\mu_0 + e_i(n)) : i \in I_1(\mathfrak{M}_0^{lin})\}$ for $\mu_0 \in \mathfrak{M}_0$ (with neither α^* nor C^* depending on the choice of $\mu_0 \in \mathfrak{M}_0$).

(c) Under the assumptions of Part (a) (Part (b), respectively) implying existence of a critical value $C(\alpha)$ satisfying (15) ((16), respectively), a smallest critical value, denoted by $C_\diamond(\alpha)$, satisfying (15) ((16), respectively) exists for every $0 < \alpha < 1$.²⁹ And $C_\diamond(\alpha)$ corresponding to Part (a) (Part (b), respectively) is also the smallest among the critical values leading to equality in (15) ((16), respectively) whenever such a critical value exists. [Although $C_\diamond(\alpha)$ corresponding to Part (a) and (b), respectively, will typically be different, we use the same symbol.]

We see from the theorem that \tilde{T}_{uc} is always size controllable over \mathfrak{C}_{Het} , but as discussed in Subsection 5.1 below there is a caveat: Unless (11), i.e., the necessary and sufficient condition for size-controllability of T_{uc} , is satisfied, size-controlling \tilde{T}_{uc} leads to trivial tests. We also see that the condition for size control of \tilde{T}_{Het} over \mathfrak{C}_{Het} , i.e., condition (17) is always satisfied in case $\tilde{\mathbf{B}} = \mathfrak{M}_0$ (since (17) is then equivalent to $e_i(n) \notin \mathfrak{M}_0^{lin}$ for every $i \in I_1(\mathfrak{M}_0^{lin})$). Furthermore, condition (17) always only depends on X and R ; in particular, it does not depend on how the weights \tilde{d}_i figuring in the definition of \tilde{T}_{Het} have been chosen (note that $\mu_0 + e_i(n) \notin \tilde{\mathbf{B}}$ is equivalent to $e_i(n) \notin \tilde{\mathbf{B}} - \mu_0$ and that the set $\tilde{\mathbf{B}} - \mu_0$ depends only on X and R). Furthermore, the size-controlling critical values $C(\alpha)$ in Part (b) of the preceding theorem depend only on X , R , and r , as well as on the choice of weights \tilde{d}_i , whereas in Part (a) the dependence is only on X , R , and r . We do not show these dependencies in the notation. In fact, as shown in Lemma D.3 in Appendix D, it turns out that the size and the size-controlling critical values in both cases actually do *not* depend on the value of r at all (provided the weights \tilde{d}_i are not allowed

²⁷Condition (17) clearly implies that the set $\tilde{\mathbf{B}}$ is a proper subset of \mathbb{R}^n and thus implies Assumption 2. Hence, we could have dropped this assumption from the formulation of the theorem. A similar remark applies to some of the other results given below and will not be repeated.

²⁸The case where \tilde{T}_{Het} is constant on $\mathbb{R}^n \setminus \tilde{\mathbf{B}}$ can actually occur under Assumption 2, see Remark D.2 in Appendix D. In such a case \tilde{T}_{Het} is trivially size-controllable (since $\tilde{\mathbf{B}}$ is a $\lambda_{\mathbb{R}^n}$ -null set under Assumption 2 and since all probability measures in (2) are absolutely continuous). However, neither a smallest size-controlling critical value exists (when considering rejection regions of the form $\{\tilde{T}_{Het} \geq C\}$) nor can exact size controllability be achieved for $0 < \alpha < 1$. [If Assumption 2 is violated, \tilde{T}_{Het} is identically zero and a similar remark applies.]

²⁹Note that there are in fact no assumptions for Part (a). We have chosen this formulations for reasons of brevity.

to depend on r in case of \tilde{T}_{Het}). Similarly, it is easy to see that α^* and C^* do not depend on r (under the same provision as before in case of \tilde{T}_{Het}).

Similarly as in Section 4, a critical value delivering size control over \mathfrak{C}_{Het} also delivers size control over *any* other heteroskedasticity model \mathfrak{C} since $\mathfrak{C} \subseteq \mathfrak{C}_{Het}$. Of course, for such a \mathfrak{C} even smaller critical values (than needed for \mathfrak{C}_{Het}) may already suffice for size control. Also note that sufficient conditions implying size control over \mathfrak{C}_{Het} may be more restrictive than sufficient conditions only implying size control over a smaller heteroskedasticity model \mathfrak{C} . For size control results tailored to such smaller models \mathfrak{C} see Appendix A.

Remark 5.2. (*Some equivalencies*) If the respective smallest size-controlling critical values are used (provided they exist), the tests obtained from \tilde{T}_{Het} with the HC0R and the HC1R weights, respectively, are identical, as these two test statistics differ only by a multiplicative constant. The same reasoning applies to the test statistics based on the HC0R-HC4R weights, respectively, in case \tilde{h}_{ii} does not depend on i .

Remark 5.3. (*Positivity of size-controlling critical values*) For every $0 < \alpha < 1$ any $C(\alpha)$ satisfying (15) or (16) is necessarily positive. To see this observe that $\{\tilde{T}_{uc} \geq C\} = \{\tilde{T}_{Het} \geq C\} = \mathbb{R}^n$ for $C \leq 0$, since both test statistics are nonnegative everywhere.

The next proposition complements Theorem 5.1 and provides a lower bound for the size-controlling critical values (other than the trivial bound given in the preceding remark). The lower bound is useful for the same reasons as discussed subsequent to Proposition 4.5.

Proposition 5.4. ^{30,31} (a) Any $C(\alpha)$ satisfying (15) necessarily has to satisfy $C(\alpha) \geq C^*$, where C^* is as in Part (a) of Theorem 5.1. In fact, for any $C < C^*$ we have $\sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(\tilde{T}_{uc} \geq C) = 1$ for every $\mu_0 \in \mathfrak{M}_0$ and every $\sigma^2 \in (0, \infty)$.

(b) Suppose Assumption 2 and (17) are satisfied, and that \tilde{T}_{Het} is not constant on $\mathbb{R}^n \setminus \tilde{\mathbf{B}}$. Then any $C(\alpha)$ satisfying (16) necessarily has to satisfy $C(\alpha) \geq C^*$, where C^* is as in Part (b) of Theorem 5.1. In fact, for any $C < C^*$ we have $\sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(\tilde{T}_{Het} \geq C) = 1$ for every $\mu_0 \in \mathfrak{M}_0$ and every $\sigma^2 \in (0, \infty)$.

Remark 5.5. Suppose the assumptions of Part (a) (Part (b), respectively) of Theorem 5.1 are satisfied. Then we know from that theorem that the size (over \mathfrak{C}_{Het}) of $\{\tilde{T}_{uc} \geq C_\diamond(\alpha)\}$ ($\{\tilde{T}_{Het} \geq C_\diamond(\alpha)\}$, respectively) equals α provided $\alpha \in (0, \alpha^*] \cap (0, 1)$. If now $\alpha^* < \alpha < 1$, then the size (over \mathfrak{C}_{Het}) of $\{\tilde{T}_{uc} \geq C_\diamond(\alpha)\}$ ($\{\tilde{T}_{Het} \geq C_\diamond(\alpha)\}$, respectively) equals α^* (where the $C_\diamond(\alpha)$'s pertaining to Parts (a) and (b) may be different). This follows from $C_\diamond(\alpha) \geq C^*$

³⁰It is not difficult to show in the context of Parts (a) and (b) of the proposition that any critical value $C > C^*$ actually leads to size less than 1. This follows from a reasoning similar as in Remark 5.4 of Pötscher and Preinerstorfer (2018).

³¹If (17) in Part (b) of the proposition does not hold, the conclusion of Part (b) can be shown to continue to hold with C^* as defined in Theorem 5.1(b), and also with C^* as defined in Lemma 5.11 of Pötscher and Preinerstorfer (2018) (note that under the assumptions of Part (b) of the proposition both definitions of C^* actually coincide as shown in the proof of Theorem 5.1). If \tilde{T}_{Het} is constant on $\mathbb{R}^n \setminus \tilde{\mathbf{B}}$ or if Assumption 2 fails (the latter implying $\tilde{T}_{Het} \equiv 0$), the conclusion of Part (b) also holds as is easily seen (regardless of which of the two definitions of C^* is adopted).

(see Proposition 5.4 above) and Remark 5.13(i) in Pötscher and Preinerstorfer (2018)).³² This argument actually also delivers that $C_{\diamond}(\alpha) = C^*$ must hold in case $\alpha^* < \alpha < 1$.

Remark 5.6. In contrast to Section 4, we have little information on the extent to which the sufficient conditions for size control in Part (b) of Theorem 5.1 are also necessary. This is due to the fact that \tilde{T}_{Het} is not a nonsphericity-corrected F-type test as noted in Remark 3.5. What can be said in general in the context of Part (b) of Theorem 5.1 in case (17) is violated, is that the size of the rejection region $\{\tilde{T}_{Het} \geq C\}$ over \mathfrak{C}_{Het} is certainly equal to 1 for every $C < \max\{\tilde{T}_{Het}(\mu_0 + e_i(n)) : e_i(n) \notin \tilde{\mathfrak{B}}\}$, where we use the convention that this maximum is $-\infty$ in case the set over which the maximum is taken is empty. This follows from Lemma 4.1 in Pötscher and Preinerstorfer (2019) with \mathbb{K} equal to the collection $\{\Pi_{(\mathfrak{M}_0^{tin})^\perp} e_i(n) : e_i(n) \notin \tilde{\mathfrak{B}}\}$.

Remark 5.7. Let \tilde{T} stand for either \tilde{T}_{Het} or \tilde{T}_{uc} in the following where in case of $\tilde{T} = \tilde{T}_{Het}$ we suppose that Assumption 2 is satisfied and that \tilde{T}_{Het} is not constant on $\mathbb{R}^n \setminus \tilde{\mathfrak{B}}$. By Lemma D.1 in Appendix D the rejection regions $\{y : \tilde{T}(y) \geq C\}$ and $\{y : \tilde{T}(y) > C\}$ differ only by a $\lambda_{\mathbb{R}^n}$ -null set. Since the measures $P_{\mu, \sigma^2 \Sigma}$ are absolutely continuous w.r.t. $\lambda_{\mathbb{R}^n}$ when Σ is nonsingular, $P_{\mu, \sigma^2 \Sigma}(\tilde{T} \geq C) = P_{\mu, \sigma^2 \Sigma}(\tilde{T} > C)$ then follows, and hence the results in this section given for rejection probabilities $P_{\mu, \sigma^2 \Sigma}(\tilde{T} \geq C)$ apply to rejection probabilities $P_{\mu, \sigma^2 \Sigma}(\tilde{T} > C)$ equally well. A similar remark applies to the results in Appendix A.2.

5.1 Tests obtained from \tilde{T}_{uc} or \tilde{T}_{Het} can be trivial

For the test statistic T_{uc} the rejection regions $\{T_{uc} \geq C\}$, as well as their complements, have positive (n -dimensional) Lebesgue measure for every positive real number C .³³ This follows from Parts 5&6 of Lemma 5.15 in Preinerstorfer and Pötscher (2016) together with Remark C.1 in Appendix C. As a consequence, all rejection probabilities – under the null as well as under the alternative – are positive and less than one regardless of the choice of $C > 0$. [This is so because of our Gaussianity assumption and the fact that all $\Sigma \in \mathfrak{C}_{Het}$ are positive definite.] For similar reasons, the same is true for T_{Het} provided Assumption 1 is satisfied.³⁴ The situation is somewhat different for tests derived from \tilde{T}_{uc} or \tilde{T}_{Het} as we shall discuss next. In the course of this, we also establish a connection between T_{uc} and \tilde{T}_{uc} that is of independent interest.

³²The assumptions for Part A of Proposition 5.12 in Pötscher and Preinerstorfer (2018) required in Remark 5.13 of that paper are satisfied under the assumptions of Theorem 5.1 as shown in the proof of Theorem A.5 in Appendix D. In this proof also the condition $\lambda_{\mathbb{R}^n}(\tilde{T}_{uc} = C^*) = 0$ ($\lambda_{\mathbb{R}^n}(\tilde{T}_{Het} = C^*) = 0$, respectively) required in Remark 5.13 of Pötscher and Preinerstorfer (2018) is verified.

³³The case $C \leq 0$ is uninteresting as the rejection region of T_{uc} (and of all other test statistics considered) then are the entire space \mathbb{R}^n , since T_{uc} (and the other test statistics considered) take on only nonnegative values.

³⁴If Assumption 1 is not satisfied then $T_{Het} \equiv 0$, and the resulting test (with rejection region $\{T_{Het} \geq C\}$) is trivial as it never rejects for $C > 0$, while it always rejects for $C \leq 0$.

5.1.1 The case of \tilde{T}_{uc}

First, observe that $\tilde{T}_{uc}(y) \leq n - (k - q)$ holds for every $y \in \mathbb{R}^n$ and that this bound is sharp. To see this, note that using standard least-squares theory

$$\tilde{T}_{uc}(y) = (n - (k - q)) \left(1 - \frac{\sum_{i=1}^n \hat{u}_i^2(y)}{\sum_{i=1}^n \tilde{u}_i^2(y)} \right) \leq n - (k - q) \quad (18)$$

for $y \notin \mathfrak{M}_0$ and that $\tilde{T}_{uc}(y) = 0$ else; the bound is attained precisely for $y \in \text{span}(X) \setminus \mathfrak{M}_0$. An immediate consequence of this observation is that any critical value $C \geq (n - (k - q))$ leads to a test with rejection region $\{T_{uc} \geq C\}$ that is either empty (if $C > n - (k - q)$) or is a $\lambda_{\mathbb{R}^n}$ -null set, namely $\text{span}(X) \setminus \mathfrak{M}_0$ (if $C = n - (k - q)$). Consequently, such a test is trivial in that all rejection probabilities (under the null as well as under the alternative) are zero (because of our Gaussianity assumption and the fact that all $\Sigma \in \mathfrak{C}_{Het}$ are positive definite). As an aside we note that any $C < n - (k - q)$ leads to a non-trivial test as is easily seen.

Of course, a critical value C satisfying $C \geq n - (k - q)$ is certainly size-controlling, but is useless since it leads to a trivial test as just discussed. We now ask if and when the smallest size-controlling critical value $C_\diamond(\alpha)$, guaranteed to exist by Part (c) of Theorem 5.1, leads to a non-trivial test. [This is certainly so if α^* in Part (a) of Theorem 5.1 is positive, but note that the theorem is silent on this issue.] To obtain insight, we establish a simple, but important, relationship between the test statistics \tilde{T}_{uc} and T_{uc} that is of independent interest also: Note that standard least-squares theory gives

$$T_{uc}(y) = (n - k) \left(\frac{\sum_{i=1}^n \tilde{u}_i^2(y)}{\sum_{i=1}^n \hat{u}_i^2(y)} - 1 \right)$$

for $y \notin \text{span}(X)$, and recall $T_{uc}(y) = 0$ for $y \in \text{span}(X)$. Hence, we obtain

$$\tilde{T}_{uc}(y) = (n - (k - q)) (T_{uc}(y) / (n - k + T_{uc}(y))) = g(T_{uc}(y)) \quad (19)$$

for every $y \notin \text{span}(X)$, where $g : [0, \infty) \rightarrow [0, n - (k - q))$ is continuous and strictly increasing with $\lim_{x \rightarrow \infty} g(x) = (n - (k - q))$. [Since $T_{uc}(y_m) \rightarrow \infty$ for every sequence $y_m \rightarrow y \in \text{span}(X) \setminus \mathfrak{M}_0$, the sharpness of the bound $n - (k - q)$ can thus also be read-off from (19).] As a consequence, for every critical value $C > 0$, the rejection regions $\{\tilde{T}_{uc} \geq C\}$ and $\{T_{uc} \geq g^{-1}(C)\}$ differ at most by $\text{span}(X)$, which is a $\lambda_{\mathbb{R}^n}$ -null set; in particular, the rejection probabilities (under the null as well as under the alternative) are the same.³⁵ *That is, the test statistics \tilde{T}_{uc} and T_{uc} give rise to (essentially) the same test, if the critical values chosen are linked by the function g as above. In particular, as we shall see, this is the case if the respective smallest size-controlling critical values are used for both test statistics (provided these exist).*

To see what the preceding discussion entails for the existence of non-trivial size-controlling critical values for \tilde{T}_{uc} we distinguish two cases. In the first case we shall see that non-trivial

³⁵This is so because of our Gaussianity assumption and the fact that all $\Sigma \in \mathfrak{C}_{Het}$ are positive definite.

size-controlling critical values do not exist, whereas in the second case they do indeed exist.

Case 1: Condition (11) is violated. Recall from Proposition 4.7 that then the size of $\{T_{uc} \geq D\}$ is 1 for every real D (in particular, implying that T_{uc} is not size controllable).³⁶ It transpires from the preceding discussion, that hence the size of $\{\tilde{T}_{uc} \geq C\}$ must equal 1 for every C satisfying $0 < C < n - (k - q)$ (and a fortiori for $C \leq 0$), because $D := g^{-1}(C)$ is well-defined and real for $0 < C < n - (k - q)$. As a consequence, any size-controlling critical value C for \tilde{T}_{uc} must satisfy $C \geq n - (k - q)$ (with the smallest size-controlling critical value given by $n - (k - q)$), thus leading to a rejection region that is trivial in that it is empty (if $C > n - (k - q)$) or is a $\lambda_{\mathbb{R}^n}$ -null set, namely $\text{span}(X) \setminus \mathfrak{M}_0$ (if $C = n - (k - q)$). That is – while \tilde{T}_{uc} is size-controllable in the present case – it is so only in a trivial way.³⁷ [Another way of arriving at the above conclusion is to use Part (a) of Proposition 5.4 and to observe that in Part (a) of Theorem 5.1 the quantity C^* equals $n - (k - q)$. To see the latter, note that violation of condition (11) implies existence of an index $i \in I_1(\mathfrak{M}_0^{lin})$ with $e_i(n) \in \text{span}(X)$. In particular, $\hat{u}(\mu_0 + e_i(n)) = 0$. Since $e_i(n) \notin \mathfrak{M}_0^{lin}$ must hold in view of $i \in I_1(\mathfrak{M}_0^{lin})$, and thus $\mu_0 + e_i(n) \notin \mathfrak{M}_0$ for every $\mu_0 \in \mathfrak{M}_0$ must be true, we may use (18) to arrive at $\tilde{T}_{uc}(\mu_0 + e_i(n)) = n - (k - q)$ for this $i \in I_1(\mathfrak{M}_0^{lin})$. This shows $C^* \geq n - (k - q)$. Equality then follows since $C^* \leq n - (k - q)$ trivially holds by (18). As a point of interest we also note that $C^* = n - (k - q)$ implies that α^* in Part (a) of Theorem 5.1 satisfies $\alpha^* = 0$.]

Case 2: Condition (11) is satisfied. In this case T_{uc} is size controllable according to Theorem 4.1. In particular, for any given $\alpha \in (0, 1)$ there exists a smallest real number $D_\diamond(\alpha)$ such that the size of $\{T_{uc} \geq D_\diamond(\alpha)\}$ is less than or equal to α , with equality holding for $\alpha \in (0, \alpha_{T_{uc}}^*] \cap (0, 1)$ where $\alpha_{T_{uc}}^*$ refers to α^* appearing in Theorem 4.1(a) and recall from that theorem that $\alpha_{T_{uc}}^* > 0$; and $D_\diamond(\alpha) > 0$ by Remark 4.4.³⁸ Also note that the rejection region $\{T_{uc} \geq D_\diamond(\alpha)\}$ is not trivial as it has positive $\lambda_{\mathbb{R}^n}$ -measure (and the same is true for its complement); see the discussion at the very beginning of this section. Setting $C_\diamond(\alpha) = g(D_\diamond(\alpha))$ and using that $\{\tilde{T}_{uc} \geq C_\diamond(\alpha)\}$ and $\{T_{uc} \geq g^{-1}(C_\diamond(\alpha))\} = \{T_{uc} \geq D_\diamond(\alpha)\}$ differ at most by the $\lambda_{\mathbb{R}^n}$ -null set $\text{span}(X)$, we see that (i) $0 < C_\diamond(\alpha) < n - (k - q)$, (ii) the size of $\{\tilde{T}_{uc} \geq C_\diamond(\alpha)\}$ is less than or equal to α , with equality holding for $\alpha \in (0, \alpha_{\tilde{T}_{uc}}^*] \cap (0, 1)$, (iii) $C_\diamond(\alpha)$ is the smallest size-controlling critical value (recall that g is strictly increasing), and (iv) the rejection region $\{\tilde{T}_{uc} \geq C_\diamond(\alpha)\}$ is not trivial as it has positive $\lambda_{\mathbb{R}^n}$ -measure (and the same is true for its complement). In particular, note that \tilde{T}_{uc} and T_{uc} give rise to (essentially) the same test if the respective smallest size-controlling critical values are used. We furthermore note that in the present situation $C_{\tilde{T}_{uc}}^* = g(C_{T_{uc}}^*)$ and $\alpha_{\tilde{T}_{uc}}^* = \alpha_{T_{uc}}^*$ hold, where $C_{T_{uc}}^*$, $\alpha_{T_{uc}}^*$ correspond to C^* , α^* in Part (a) of Theorem 4.1, whereas $C_{\tilde{T}_{uc}}^*$, $\alpha_{\tilde{T}_{uc}}^*$ correspond to C^* , α^* in Part (a) of Theorem 5.1.³⁹ In particular, $\alpha_{\tilde{T}_{uc}}^* > 0$ and $0 \leq C_{\tilde{T}_{uc}}^* < n - (k - q)$ follow. These claims can be seen as follows: Under condition

³⁶Recall that in this section (i.e., Section 5) size always refers to size over \mathfrak{C}_{Het} .

³⁷The trivial size-controlling critical values C for \tilde{T}_{uc} sort of correspond to using ∞ as a “size-controlling critical value” for T_{uc} .

³⁸If $\alpha_{T_{uc}}^* < \alpha < 1$, then the size, in fact, equals $\alpha_{T_{uc}}^*$; see Remark 4.6.

³⁹If $\alpha_{T_{uc}}^* < \alpha < 1$, then the size of $\{\tilde{T}_{uc} \geq C_\diamond(\alpha)\}$ is, in fact, equal to $\alpha_{T_{uc}}^* = \alpha_{\tilde{T}_{uc}}^*$; cf. Footnote 38 and Remark 5.5.

(11) we have $\mu_0 + e_i(n) \notin \text{span}(X)$ for every $i \in I_1(\mathfrak{M}_0^{lin})$ and every $\mu_0 \in \mathfrak{M}_0$. Consequently, $\tilde{T}_{uc}(\mu_0 + e_i(n)) = g(T_{uc}(\mu_0 + e_i(n)))$, which proves $C_{\tilde{T}_{uc}}^* = g(C_{T_{uc}}^*)$ in view of strict monotonicity of g . The relation $\alpha_{\tilde{T}_{uc}}^* = \alpha_{T_{uc}}^*$ then follows from the definitions of $\alpha_{\tilde{T}_{uc}}^*$ and $\alpha_{T_{uc}}^*$ using that $\{\tilde{T}_{uc} \geq C\}$ and $\{T_{uc} \geq g^{-1}(C)\}$ differ at most by the $\lambda_{\mathbb{R}^n}$ -null set $\text{span}(X)$ for every $C > 0$. Positivity of $\alpha_{\tilde{T}_{uc}}^*$ now follows from positivity of $\alpha_{T_{uc}}^*$ discussed before, and $C_{\tilde{T}_{uc}}^* < n - (k - q)$ follows since $C_{\tilde{T}_{uc}}^* = g(C_{T_{uc}}^*)$ and $C_{T_{uc}}^* < \infty$. [Another way of proving $\alpha_{\tilde{T}_{uc}}^* > 0$ and $0 \leq C_{\tilde{T}_{uc}}^* < n - (k - q)$ without using relationship (19), is to first establish $C_{\tilde{T}_{uc}}^* < n - (k - q)$ (from observing that $\hat{u}(\mu_0 + e_i(n)) \neq 0$ (as $\mu_0 + e_i(n) \notin \text{span}(X)$) for every $i \in I_1(\mathfrak{M}_0^{lin})$, which implies $\tilde{T}_{uc}(\mu_0 + e_i(n)) < n - (k - q)$ for every such i in view of (18)) and then to proceed analogously as in the proof of Theorem 5.8 below.]

While \tilde{T}_{uc} is always size-controllable, whereas T_{uc} is not, this does not represent any real advantage of \tilde{T}_{uc} over T_{uc} , as we have seen that \tilde{T}_{uc} admits only trivial size-controlling critical values in the case where T_{uc} is not size-controllable. Even more importantly, and already noted above, these test statistics give rise to (essentially) the same test if for both test statistics the respective smallest size-controlling critical values are used (provided they exist).

5.1.2 The case of \tilde{T}_{Het}

For \tilde{T}_{Het} we find that, not infrequently, it is also a bounded function, although we have no proof that this is always so. We illustrate the problems that can arise here first by an example.

Example 5.1. Consider the $n \times 2$ design matrix X where the first column represents an intercept, the second column is $x := (1, -1, 0, \dots, 0)'$, and $n \geq 3$. Let $R = (0, 1)$, $r = 0$, hence $q = 1$. Obviously, the first column of X spans \mathfrak{M}_0^{lin} . Since $e_i(n) \notin \mathfrak{M}_0^{lin}$ for every $i = 1, \dots, n$, Assumption 2 holds. Furthermore, $\tilde{h}_{ii} = n^{-1}$. Thus $\tilde{d}_i = \tilde{d}_1$ holds for every $i = 1, \dots, n$ and for every of the five choices HC0R-HC4R. Note that $\tilde{d}_1^{-1} = 1$ (HC0R), $\tilde{d}_1^{-1} = 1 - n^{-1}$ (HC1R), $\tilde{d}_1^{-1} = 1 - n^{-1}$ (HC2R), $\tilde{d}_1^{-1} = (1 - n^{-1})^2$ (HC3R), and $\tilde{d}_1^{-1} = 1 - n^{-1}$ (HC4R), and hence $0 < \tilde{d}_1^{-1} \leq 1$ for all five choices. Straightforward computations now show that $\tilde{\Omega}_{Het}(y) = \tilde{d}_1 \left[(y_1 - \bar{y})^2 + (y_2 - \bar{y})^2 \right] / 4$ and

$$\tilde{T}_{Het}(y) = \tilde{d}_1^{-1} (y_1 - y_2)^2 / \left[(y_1 - \bar{y})^2 + (y_2 - \bar{y})^2 \right] \quad (20)$$

whenever the numerator is positive, and $\tilde{T}_{Het}(y) = 0$ otherwise. Here \bar{y} denotes the arithmetic mean of the observations y_i . [For later use we also note that the set $\tilde{\mathbf{B}}$ is given by $\{y \in \mathbb{R}^n : y_1 = y_2 = \bar{y}\}$, and that the size control condition (17) is satisfied, since $e_i(n) \notin \tilde{\mathbf{B}}$ for every $i = 1, \dots, n$ (also note that μ_0 can be chosen to be zero because of $r = 0$). Furthermore, \tilde{T}_{Het} is not constant on $\mathbb{R}^n \setminus \tilde{\mathbf{B}}$, since $\tilde{T}_{Het}(e_1(n)) = \tilde{T}_{Het}(e_2(n)) = \tilde{d}_1^{-1} n^2 / [(n-1)^2 + 1]$ and $\tilde{T}_{Het}(e_i(n)) = 0$ for $i \geq 3$ (note $n \geq 3$) and since $e_i(n) \notin \tilde{\mathbf{B}}$ for every i .] It is now evident from (20) that $\tilde{T}_{Het}(y) \leq 2\tilde{d}_1^{-1}$ for every $y \in \mathbb{R}^n$ and that this bound is attained whenever $y_1 + y_2 = 2\bar{y}$ (e.g., for $y = x$). It follows that any critical value $C \geq 2\tilde{d}_1^{-1}$ leads to a test with rejection region that is empty if $C > 2\tilde{d}_1^{-1}$, and is a Lebesgue null-set if $C = 2\tilde{d}_1^{-1}$ (the latter following from Lemma D.1(d) in

Appendix D together with some of the observations just noted after (20)); thus in both cases all the rejection probabilities are zero under the null as well as under the alternative (given our Gaussianity assumption and the fact that all $\Sigma \in \mathfrak{C}_{Het}$ are positive definite); in particular, these tests have zero power. Since $\tilde{d}_1^{-1} \leq 1$, this eliminates all critical values $C \geq 2$ from practical use. In particular, this eliminates the commonly used choice where C is the 95%-quantile of a chi-square distribution with 1 degree of freedom, which is approximately equal to 3.8415.

In the preceding example any critical value $C \geq 2\tilde{d}_1^{-1}$ is trivially a size-controlling critical value for the given significance level α ($0 < \alpha < 1$), but it is “too large” and leads to a trivial test. Certainly, one would prefer to use the smallest size-controlling critical value $C_\diamond(\alpha)$ instead (which in the preceding example exists by Theorem 5.1 and by what has been shown in the example) and one would hope that the resulting test is not trivial. As we shall show, this is indeed the case. To this end we first give a general result that, in particular, is applicable to the preceding example. Recall that $C_\diamond(\alpha)$ is positive (Remark 5.3), and that Theorem 5.1 is silent on whether $\alpha^* > 0$ or not.

Theorem 5.8. *Suppose Assumption 2 and (17) are satisfied, and that \tilde{T}_{Het} is not constant on $\mathbb{R}^n \setminus \tilde{\mathbf{B}}$. Let α satisfy $0 < \alpha < 1$, and let C^* and α^* be as defined in Part (b) of Theorem 5.1. If $C^* < \sup_{y \in \mathbb{R}^n} \tilde{T}_{Het}(y)$ holds, then we have $\alpha^* > 0$, and the rejection region $\{\tilde{T}_{Het} \geq C_\diamond(\alpha)\}$ is not a $\lambda_{\mathbb{R}^n}$ -null set, where $C_\diamond(\alpha)$ is the smallest size-controlling critical value as in Part (c) of Theorem 5.1.*

Remark 5.9. (i) The preceding theorem clearly implies that – under its assumptions – the rejection probabilities associated with the rejection region $\{\tilde{T}_{Het} \geq C_\diamond(\alpha)\}$ are positive under the null as well as under the alternative (in view of our Gaussianity assumption and the fact that all $\Sigma \in \mathfrak{C}_{Het}$ are positive definite). [While we already know from Theorem 5.1(b) and Remark 5.5 that the rejection region $\{\tilde{T}_{Het} \geq C_\diamond(\alpha)\}$ has size equal to α in case $\alpha \in (0, \alpha^*] \cap (0, 1)$, and has size equal to α^* if $\alpha^* < \alpha < 1$, this by itself does *not* allow one to conclude that the rejection region has positive $\lambda_{\mathbb{R}^n}$ -measure as the case $\alpha^* = 0$ is not ruled out by Theorem 5.1(b) and Remark 5.5.]

(ii) Suppose $C^* = \sup_{y \in \mathbb{R}^n} \tilde{T}_{Het}(y)$, but that the other assumptions of Theorem 5.8 hold.⁴⁰ Then the rejection region $\{\tilde{T}_{Het} \geq C_\diamond(\alpha)\}$ is a $\lambda_{\mathbb{R}^n}$ -null set; thus also the smallest (and hence any) size-controlling critical value leads to a trivial test. To prove the claim, note that by Proposition 5.4 we have $C_\diamond(\alpha) \geq C^*$, implying that the rejection regions are either empty or coincide with the sets $\{\tilde{T}_{Het} = C^*\}$, respectively. In the latter case apply Part (d) of Lemma D.1 in Appendix D. We also point out that in the present case $\alpha^* = 0$ must hold since the rejection regions appearing in the definition of α^* are all empty (because of $C > C^* = \sup_{y \in \mathbb{R}^n} \tilde{T}_{Het}(y)$ in the definition of α^*).

Example 5.2. We continue the discussion of Example 5.1. As noted prior to Theorem 5.8, any

⁴⁰We have not investigated whether this case can actually occur for \tilde{T}_{Het} . Recall that for \tilde{T}_{uc} this case indeed can occur, see Case 1 in Section 5.1.1.

critical value $C \geq 2\tilde{d}_1^{-1}$ is size-controlling in a trivial way, but leads to trivial rejection regions. We now show that the smallest size-controlling critical value $C_\diamond(\alpha)$ indeed leads to a non-trivial test (which, in particular, has positive rejection probabilities in view of our Gaussianity assumption and the fact that all $\Sigma \in \mathfrak{C}_{Het}$ are positive definite). For this it suffices to verify the assumptions of Theorem 5.8. The first three assumptions have already been verified above. From the calculations in Example 5.1 it is now easy to see that $C^* = \tilde{d}_1^{-1}n^2/[(n-1)^2 + 1]$, which is smaller than $2\tilde{d}_1^{-1} = \sup_{y \in \mathbb{R}^n} \tilde{T}_{Het}(y)$. This completes the proof of the assertion. From Remark 5.9(i) we furthermore see that the rejection region $\{\tilde{T}_{Het} \geq C_\diamond(\alpha)\}$ has size equal to α if $\alpha \in (0, \alpha^*] \cap (0, 1)$, and has size equal to α^* if $\alpha^* < \alpha < 1$. Finally we note that size-controlling critical values that do not lead to trivial tests must lie in the interval $[\tilde{d}_1^{-1}n^2/[(n-1)^2 + 1], 2\tilde{d}_1^{-1})$ which is quite narrow as it is contained in the interval $[\tilde{d}_1^{-1}, 2\tilde{d}_1^{-1})$.

While the situation in Example 5.1 is somewhat particular, the example may perhaps contribute to a better understanding of the Monte Carlo findings in Davidson and MacKinnon (1985) and Godfrey (2006), namely that the tests, obtained from \tilde{T}_{Het} (employing HC0R-HC4R weights) in conjunction with conventional critical values such as the 95%-quantile of a chi-square distribution with appropriate degrees of freedom, can suffer from severe underrejection under the null.

6 Generalizations

6.1 Generalizations beyond Gaussianity

(i) All results in the preceding sections (as well as the extensions described in Appendix A) referring to properties under the null hypothesis carry over as they stand to the situation where the error term \mathbf{U} in (1) is elliptically symmetric distributed and has no atom at zero, i.e., \mathbf{U} is distributed as $\sigma\Sigma^{1/2}\mathbf{z}$ where \mathbf{z} has a spherically symmetric distribution on \mathbb{R}^n that has no atom at zero.⁴¹ This is so since – under this distributional model – the null rejection probabilities of any $G(\mathfrak{M}_0)$ -invariant rejection region coincide with the corresponding null rejection probabilities under the Gaussian model (i.e., where \mathbf{z} is standard Gaussian); see the discussion in Section 5.5 of Preinerstorfer and Pötscher (2016) and Appendix E.1 of Pötscher and Preinerstorfer (2018).⁴² This implies, in particular, not only that the sufficient conditions for size controllability under the above elliptically symmetric distributed model as well as under the Gaussian model are the same, but that also the numerical values of the size-controlling critical values coincide. As a consequence, the algorithms for computing the size-controlling critical values in the Gaussian case (used in Section 9 and described in Appendix E) can be used in the above elliptically symmetric distributed case without any change whatsoever. The same is actually true if \mathbf{z} has a

⁴¹Note that all results in the preceding sections, except for a few comments in Section 5.1, are results referring to properties under the null hypothesis,

⁴²Note that all rejection regions considered in the preceding sections are $G(\mathfrak{M}_0)$ -invariant, because the test statistics considered are so.

distribution in a certain class larger than the class of spherical symmetric distributions with no atom at zero, see Appendix E.1 of Pötscher and Preinerstorfer (2018).

(ii) Furthermore, as discussed in detail in Appendix E.2 of Pötscher and Preinerstorfer (2018), the sufficient conditions for size controllability that we have derived under Gaussianity also imply size controllability for many more forms of distribution of \mathbf{z} than those mentioned in (i); however, the corresponding size-controlling critical values may then differ from the size-controlling critical values that apply under Gaussianity.

(iii) Similarly as in Section 5.5 of Preinerstorfer and Pötscher (2016), the negative results given in the preceding sections (as well as the ones described in Appendix A) such as, e.g., size 1 results, extend in a trivial way beyond the Gaussian model as long as the maintained assumptions on the feasible error distributions are weak enough to ensure that the implied (possibly semiparametric) model, i.e., set of distributions for \mathbf{Y} , contains the set given in (2), but possibly contains also other distributions.

(iv) A further generalization beyond Gaussianity in the important special case where $\mathfrak{C} = \mathfrak{C}_{Het}$ is as follows: Suppose \mathbf{U} is distributed as $\sigma \Sigma^{1/2} \text{diag}(\mathbf{r})\mathbf{z}$ where \mathbf{z} is standard normally distributed on \mathbb{R}^n and where the n -dimensional random vector \mathbf{r} is independent of \mathbf{z} with distribution ρ , where ρ is a distribution on $(0, \infty)^n$. [This includes the case where the elements of $\text{diag}(\mathbf{r})\mathbf{z}$ form an i.i.d. sample from a scale mixture of normals.] Let $Q_{\mu, \sigma^2 \Sigma, \rho}$ denote the implied distribution for \mathbf{Y} given by (1) where $\mu = X\beta$. Consider now instead of (2) the (semiparametric) model given by all distributions $Q_{\mu, \sigma^2 \Sigma, \rho}$ where $\mu \in \text{span}(X)$, $0 < \sigma^2 < \infty$, $\Sigma \in \mathfrak{C}$, and ρ is an arbitrary distribution on $(0, \infty)^n$. Then the sufficient conditions for size controllability derived under Gaussianity in earlier sections (and in Appendix A) also imply size controllability in this larger model. In fact, the size-controlling critical values that apply under Gaussianity deliver also size control under this more general model. This follows from the following reasoning: Let W be a Borel set in \mathbb{R}^n such that $P_{\mu_0, \sigma^2 \Sigma}(W) \leq \alpha$ for every $\mu_0 \in \mathfrak{M}_0$, every $0 < \sigma^2 < \infty$, and every $\Sigma \in \mathfrak{C}_{Het}$. Then for every such μ_0 , σ^2 , Σ , and every distribution ρ on $(0, \infty)^n$ we have

$$\begin{aligned} Q_{\mu_0, \sigma^2 \Sigma, \rho}(W) &= \Pr(\mu_0 + \sigma \Sigma^{1/2} \text{diag}(\mathbf{r})\mathbf{z} \in W) = \mathbb{E}[\Pr(\mu_0 + \sigma \Sigma^{1/2} \text{diag}(\mathbf{r})\mathbf{z} \in W | \mathbf{r})] \\ &= \mathbb{E}[\Pr(\mu_0 + \sigma_{\mathbf{r}} \Sigma_{\mathbf{r}}^{1/2} \mathbf{z} \in W | \mathbf{r})] = \mathbb{E}[P_{\mu_0, \sigma_{\mathbf{r}}^2 \Sigma_{\mathbf{r}}}(W)] \leq \alpha, \end{aligned}$$

where $\Sigma_{\mathbf{r}}^{1/2} := \Sigma^{1/2} \text{diag}(\mathbf{r})/s_{\mathbf{r}}$ with $s_{\mathbf{r}}$ denoting the positive square root of the sum of the diagonal elements of $(\Sigma^{1/2} \text{diag}(\mathbf{r}))^2 = \Sigma \text{diag}^2(\mathbf{r})$ and where $\sigma_{\mathbf{r}} = \sigma s_{\mathbf{r}}$. Here we have used that $P_{\mu, \sigma_{\mathbf{r}}^2 \Sigma_{\mathbf{r}}}(W) \leq \alpha$ by assumption since $\Sigma_{\mathbf{r}} = \Sigma \text{diag}^2(\mathbf{r})/s_{\mathbf{r}}^2 \in \mathfrak{C}_{Het}$ and $0 < \sigma_{\mathbf{r}} < \infty$ hold for every realization of \mathbf{r} . In the above \Pr denotes the probability measure governing (\mathbf{r}, \mathbf{z}) and \mathbb{E} the corresponding expectation operator. [In the special case where $\text{diag}(\mathbf{r})$ is a (random) multiple of the identity matrix I_n , the assumption $\mathfrak{C} = \mathfrak{C}_{Het}$ is superfluous as then $\Sigma_{\mathbf{r}}^{1/2} = \Sigma^{1/2}$, which by assumption belongs to the given \mathfrak{C} . In this case \mathbf{U} satisfies the assumptions in (i), and hence (iv) adds little new, except that – in contrast to (i) – the reasoning works without use of $G(\mathfrak{M}_0)$ -invariance.]

(v) It is apparent from the reasoning in (iv) that Gaussianity of \mathbf{z} can be replaced by any other distributional assumption for which size controllability has already been established. E.g., one can in (iv) choose \mathbf{z} to have a spherically symmetric distribution without an atom at zero or to have a distribution in the more general class mentioned in (i) (note that all relevant rejection regions discussed in earlier sections are $G(\mathfrak{M}_0)$ -invariant and thus (i) applies). In a similar vein, one can combine the results in Appendix E.2 of Pötscher and Preinerstorfer (2018) discussed in (ii) above with the reasoning outlined in (iv). We abstain from presenting details.

6.2 Generalizations to stochastic regressors

The assumption of nonstochastic regressors can be easily relaxed as follows: If X is random and \mathbf{U} is conditionally on X distributed as $N(0, \sigma^2 \Sigma)$, with $\sigma^2 = \sigma^2(X) > 0$ and $\Sigma = \Sigma(X) \in \mathfrak{C} \subseteq \mathfrak{C}_{Het}$, the results of the paper obviously apply after one conditions on X provided almost all realizations of X satisfy the assumptions of our theorems, which will typically be the case (for brevity we do not provide a formal statement here). The resulting conditional size control statements then immediately imply that the so-obtained conditional size-controlling critical values $C = C(\alpha, X)$ also control size unconditionally. And again similar generalizations to non-Gaussianity as discussed in Section 6.1 are possible here.

7 Results for other classes of tests

The results in Sections 4 and 5 (and in Appendix A) have been obtained with the help of a general theory developed in Section 5 of Preinerstorfer and Pötscher (2016), Section 5 of Pötscher and Preinerstorfer (2018), and Section 3.1 of Pötscher and Preinerstorfer (2019) that covers a very broad class of test statistics (and actually allows also for correlated errors). We note that, like in Section 6.1, Gaussianity is again not essential for a good portion of this general theory, see Section 5.5 of Preinerstorfer and Pötscher (2016) as well as Appendix E of Pötscher and Preinerstorfer (2018).⁴³ We next discuss a few further situations that can also be handled by the general theory just mentioned but we refrain from spelling out the details:⁴⁴

(i) The test statistic considered is an OLS-based test statistic like T_{Het} , but where $\hat{\Omega}_{Het}$ is now replaced by an appropriate estimator derived from a given (possibly misspecified) *parametric* heteroskedasticity model described by a parameter vector θ .

(ii) The test statistic is a Wald-type test statistic based on a (feasible) generalized least-squares estimator together with an appropriate covariance matrix estimator based on a given (possibly misspecified) parametric model. [This includes the (quasi-)maximum likelihood estimator (provided θ is unrelated to β).] Alternatively, the test statistic is the (quasi-)likelihood

⁴³Also arguments like in (iv) and (v) of Section 6.1 can be applied to obtain generalizations.

⁴⁴Applying some of the main results of this general theory (e.g., Corollary 5.6 or Proposition 5.12 of Pötscher and Preinerstorfer (2018)) will require one to determine the set $\mathbb{J}(\mathcal{L}, \mathfrak{C})$ defined in Appendix B. For the important cases $\mathfrak{C} = \mathfrak{C}_{Het}$ and $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$ (defined in Appendix A), this is already accomplished in Propositions B.1 and B.2 in Appendix B below.

ratio or (quasi-)score test statistic based on this parametric model.

(iii) The test statistic is a Wald-type test statistic as in (ii), except that the covariance matrix estimator is now nonparametric (in the spirit of heteroskedasticity robust testing) as described in Romano and Wolf (2017). See also Cragg (1983, 1992), Flachaire (2005), Wooldridge (2010, 2012), Romano and Wolf (2017), Lin and Chou (2018), DiCiccio et al. (2019).

8 Some comments on power

Under our maintained assumptions, heteroskedasticity robust tests based on T_{Het} or T_{uc} (using an arbitrary critical value C , including size-controlling ones) have positive power everywhere in the alternative (cf. the discussion at the beginning of Section 5.1). These tests can furthermore be shown to have power that goes to one as one moves away from the null hypothesis along sequences $(\mu_l, \sigma_l^2, \Sigma_l)$ where μ_l moves further and further away from \mathfrak{M}_0 (the affine space of means described by the restrictions $R\beta = r$) in an orthogonal direction as $l \rightarrow \infty$, where σ_l^2 converges to some finite and positive σ^2 , and Σ_l converges to a positive definite matrix. Despite of what has just been said, these tests can have, in fact not infrequently will have, *infimal* power equal to zero if \mathfrak{C} is sufficiently rich, e.g., if $\mathfrak{C} = \mathfrak{C}_{Het}$; cf. Theorem 4.2 in Preinerstorfer and Pötscher (2016), Lemma 5.11 in Pötscher and Preinerstorfer (2018), and Theorem 4.2 in Pötscher and Preinerstorfer (2019). [This does not contradict the before mentioned result as for this result sequences Σ_l that converge to a singular matrix as $l \rightarrow \infty$ were ruled out.]

For tests based on \tilde{T}_{Het} or \tilde{T}_{uc} the situation is somewhat different. As shown in Section 5.1, tests based on \tilde{T}_{Het} or \tilde{T}_{uc} can be trivial for some choices of critical values C (and then will have power zero everywhere in the alternative). However, if C is chosen to be the smallest size-controlling critical value (provided it exists), the resulting tests obtained from \tilde{T}_{Het} or \tilde{T}_{uc} will typically have positive power (under appropriate assumptions). In particular, then the test based on \tilde{T}_{uc} has the same power function as the test based on T_{uc} that uses its smallest size-controlling critical value, provided the latter exists, see Section 5.1.1. We have not further investigated the power properties of the tests based on \tilde{T}_{Het} in any more detail on a theoretical level. The numerical results in Section 9.2 seem to suggest that for these tests power may not go to one along sequences $(\mu_l, \sigma_l^2, \Sigma_l)$ as mentioned above: in fact, power does not rise above the significance level α in some examples (on the range of alternatives considered). This feature makes tests based on \tilde{T}_{Het} rather undesirable.

9 Numerical results

In this section we pursue two goals:

1. In Subsection 9.1 we show numerically that any of the usual heteroskedasticity robust tests can suffer from overrejection of the null hypothesis (sometimes by a large margin) when they are based on conventional critical values. While this adds to similar evidence already

present in the literature for the HC0-HC4 based tests (see Section 1), this seems to be a new observation for the HC0R-HC4R based tests. In any case, this drives home the point that none of these heteroskedasticity robust tests based on conventional critical values comes with a guarantee that size is controlled by the nominal significance level α . Consequently, instead of using conventional critical values, this strongly suggests to use size-controlling critical values as investigated in this paper.

2. In Subsection 9.2 we then numerically compute size-controlling critical values and study the power behavior of tests based on such size-controlling critical values in some examples.

In this section (and in the attending Appendices E and F) we shall often refer to T_{Het} as HC0-HC4 when we want to stress that the weights d_i being used are the HC0-HC4 weights, respectively, see Section 3. Similarly, we shall refer to \tilde{T}_{Het} as HC0R-HC4R when the HC0R-HC4R weights are used. For reasons of uniformity of notation, we shall then often denote T_{uc} as UC and \tilde{T}_{uc} as UCR. Furthermore, throughout this section we consider the heteroskedastic Gaussian linear model with $\mathfrak{C} = \mathfrak{C}_{Het}$ as introduced in Section 2; in particular, the notion of size in the present section (and the attending appendices) *always* refers to this model.

The algorithms for computing rejection probabilities, the size of a test, and size-controlling critical values used in the before-mentioned numerical computations are described in Appendix E. Implementations are available as an R-package **hrt** (Preinerstorfer (2021)). As noted at the beginning of Appendix E, some of these algorithms remain valid without any need for modification for elliptically symmetric distributed data.

9.1 Tests based on conventional critical values

We consider the important case $q = 1$, and first illustrate numerically that none of the test statistics UC, HC0-HC4, UCR, and HC0R-HC4R combined with the critical value $C_{\chi^2, 0.05} \approx 3.8415$ results in a test that is *guaranteed* to have size less than or equal to $\alpha = 0.05$. This is achieved by providing instances of design matrices X and of hypotheses, described by (R, r) , such that the respective test has size larger than the nominal significance level $\alpha = 0.05$, often by a large margin. Here $C_{\chi^2, 0.05}$ denotes the 95%-quantile of a chi-square-distribution with 1 degree of freedom. [This critical value has a justification for use with HC0-HC4 or HC0R-HC4R via asymptotic considerations, but, in general, there is no such justification for use with UC or UCR, which we nevertheless include here for completeness.⁴⁵] That is, in the instances we exhibit, this conventional critical value turns out to be *too small*. We next show similar results for other suggestions of critical values, e.g., for “degree-of-freedom” adjustments to the conventional chi-square based critical value such as the Bell-McCaffrey adjustment (Bell and McCaffrey (2002), Imbens and Kolesár (2016)). It is important to note here that in all the instances mentioned our conditions for size-controllability are satisfied, showing that size-controlling critical values can

⁴⁵Of course, in the special case of homoskedasticity, the before mentioned justification also applies to UC and UCR.

actually be found; hence, the overrejection problems mentioned before are *not* intrinsic problems, but only reflect the fact that conventional critical values can be a bad choice and do not guarantee size control. [In the present context it is worth recalling that for the test statistics HC0R-HC4R we have already shown in Example 5.1 in Section 5.1 that other situations can be found in which conventional critical values such as, e.g., $C_{\chi^2,0.05}$ are *too large*, as the resulting tests reject with probability 0 only (under the null as well as under the alternative), rendering these tests useless.]

To uncover instances where the conventional critical value $C_{\chi^2,0.05}$ is too small, we make use of the following observation: In case a given test statistic from the above list (together with a given design matrix X and hypothesis described by (R, r)) is such that the lower bound C^* on size-controlling critical values obtained in Propositions 4.5 (5.4, respectively) exceeds $C_{\chi^2,0.05}$, we are done, as we then know that the critical value $C_{\chi^2,0.05}$ leads to a test that has size 1. [As noted subsequent to Theorems 4.1 and 5.1, the value of r actually plays no rôle here.]

Since the lower bounds C^* for size-controlling critical values in Propositions 4.5 (5.4, respectively) depend on the given test statistic, on X and on R , we may – for every given choice of test statistic – numerically search for particularly “hostile” design matrices, i.e., for design matrices for which the lower bound is large, to see whether matrices X exist for which the lower bound exceeds $C_{\chi^2,0.05}$. We only do this for $k = 2$, $R = (0, 1)$, and $n = 25$, and restrict ourselves to matrices X with first column representing an intercept. The concrete search used is detailed in Appendix F.1, see Algorithm 5 in particular. Table 1 provides the so-obtained worst-case lower bounds C^* for all test statistics considered. In combination with the theoretical results

UC	731.60	UCR	23.59
HC0	95.56	HC0R	1.08
HC1	1711.19	HC1R	1.04
HC2	52.23	HC2R	1.04
HC3	1.00	HC3R	1.00
HC4	1.02	HC4R	1.04

Table 1: Worst-case lower bounds.

UC	0.98	UCR	0.98
HC0	0.99	HC0R	0.16
HC1	1.00	HC1R	0.17
HC2	0.99	HC2R	0.17
HC3	0.19	HC3R	0.14
HC4	0.11	HC4R	0.10

Table 2: Worst-case sizes using $C_{\chi^2,0.05}$.

from Propositions 4.5 and 5.4, Table 1 shows that for some design matrices X the critical value $C_{\chi^2,0.05} \approx 3.8415$ results in a test with size equal to 1 when combined with UC, HC0-HC2, and also with UCR. [This is so despite the fact that for all X matrices encountered in the numerical procedure the sufficient conditions for size-control in the pertaining theorems in Sections 4 and 5 are satisfied (as we have checked), and hence it is known that size-controlling critical values exist for any of the twelve test statistics!] Table 1 is not informative about the size of the remaining seven tests, since the corresponding entries in that table are all less than $C_{\chi^2,0.05}$. To obtain insight into the sizes of the remaining seven tests we do the following: for each of the tests we numerically compute the size for various instances of design matrices (ones that have already been used in the numerical search leading to Table 1) and report the largest one of these sizes in Table 2. We actually do this for all twelve tests considered. The algorithm used in the size computation is the implementation of Algorithm 1 in the R-package **hrt** (Preinerstorfer (2021)),

cf. the description in Appendices E.2 and F.1. Table 2 now clearly shows that for *every* test statistic considered an instance can be found, in which the size of the test (when using the critical value $C_{\chi^2, 0.05}$) clearly exceeds the nominal significance level $\alpha = .05$. The lowest value in that table is attained by HC4R, but a size of 0.10 is still twice the nominal significance level α .

We note that the numbers shown in Table 2 actually only represent numerically determined lower bounds for the actual sizes, as their computation involves (for any given X) a numerical search procedure (over the set \mathfrak{C}_{Het}) for the worst-case null rejection probability; that is, the numbers shown in Table 2 correspond to the null rejection probability computed from a “bad” covariance matrix Σ , but potentially not for the “worst” possible one. [In this process, for any given $\Sigma \in \mathfrak{C}_{Het}$, we have to numerically compute the null rejection probability, which can be done quite accurately in case $q = 1$ by algorithms like the Davies algorithm, see Appendix E.1 as well as Appendix E.2.] In particular, the entries in the 0.98-0.99 range in Table 2 are numerically determined lower bounds for the size, which, in fact, we know to be equal to 1 in light of Table 1. [We could have used this knowledge to replace the entries in question in Table 2 by 1, but we decided otherwise in order to showcase the concrete outcome of the numerical algorithm that has been run. Of course, one could also improve this outcome by using a higher accuracy parameter in the optimization procedures involved.]

Sometimes – without much theoretical justification in general – it is suggested in the literature to replace $C_{\chi^2, 0.05}$ by the 95%-quantile of an $F_{1, n-k}$ -distribution, which is approximately 4.28 in the situation considered here ($n - k = 23$). Obviously, from Table 1 we see that the conclusions regarding UC, HC0-HC2, and UCR remain the same when this critical value is used. Repeating the exercise that has led to Table 2, but with $C_{\chi^2, 0.05}$ replaced by the 95%-quantile of an $F_{1, n-k}$ -distribution, gives Table 3, leading essentially to the same conclusions.

UC	0.98	UCR	0.98
HC0	0.99	HC0R	0.15
HC1	1.00	HC1R	0.16
HC2	0.98	HC2R	0.15
HC3	0.18	HC3R	0.13
HC4	0.09	HC4R	0.08

Table 3: Worst-case sizes using F-critical value.

“Degree-of-freedom” adjustments to the conventional chi-square based critical value such as the Bell-McCaffrey adjustment (Bell and McCaffrey (2002)) have been discussed in the literature. In particular, Imbens and Kolesár (2016) suggested to use this adjustment with the HC2 statistic. We have repeated the above exercise that has led to the entry for HC2 in Table 2, but with $C_{\chi^2, 0.05}$ replaced by the Bell-McCaffrey adjustment. For the computation of the Bell-McCaffrey adjustment we relied on the R-package **dfadjust** (Kolesár (2019)). For the resulting test, the largest size that was found in our computations was 0.24, which is more than four times the nominal significance level. It transpires that this adjustment does also not come with a size-guarantee.

We conclude here by stressing that the negative findings in this subsection were obtained in a very simple model with only two regressors and where only one of the parameters is subject to test. For more complex models and test problems the size distortions may even be worse.

9.2 Power comparison of tests based on size-controlling critical values

A power comparison of two tests, both conducted at a given *nominal* significance level α , makes sense only if both tests actually are level α tests, i.e., if both tests have a size not exceeding the given α . For this reason, we now compare the tests obtained from the statistics UC, HC0-HC4, UCR, HC0R-HC4R only when respective *smallest size-controlling critical values* are used. Our theoretical results concerning the existence of size-controlling critical values, together with the algorithms for their computation in Appendix E, allow for such a comparison in terms of power.

Throughout, in addition to the power functions of the before-mentioned tests, we also show as a benchmark the power function of the *infeasible* (i.e., oracle) GLS-based F -test conducted at the 5%-significance level, that makes use of knowledge of Σ . For given $\Sigma \in \mathfrak{C}_{Het}$, the distribution of this infeasible GLS-based F -test statistic is (under $P_{X\beta, \sigma^2\Sigma}$ with $\beta \in \mathbb{R}^k$, $\sigma^2 \in (0, \infty)$) a noncentral $F_{1, n-k}$ -distribution with noncentrality parameter δ^2 , where

$$\delta = (R(X'\Sigma^{-1}X)R')^{-1/2}(R\beta - r)/\sigma.$$

Since the power functions of all the tests considered in our study depend on the parameters β , σ^2 , and Σ only through $(R\beta - r)/\sigma$ and Σ (because of $G(\mathfrak{M}_0)$ -invariance and Proposition 5.4 in Preinerstorfer and Pötscher (2016)), and thus depend only on δ and Σ , we shall – for given Σ – present all these power functions as a function of δ . We show only results for $\delta \geq 0$, as the power functions in fact depend on δ only through $|\delta|$ (for given Σ); see Proposition 5.4 in Preinerstorfer and Pötscher (2016).

9.2.1 Comparing the means of two heteroskedastic groups

As a practically relevant example, we here consider comparing the power of tests based on size-controlling critical values in the context of Example 4.4. That is, we treat the problem of comparing the means of two heteroskedastic groups (e.g., a treatment and a control group), the null hypothesis being that the difference of expected outcomes in each group is zero. We consider the case where $n = 30$ and $\alpha = 0.05$. Furthermore, we vary the size n_1 of the first group ($n_1 \in \{3, 9, 15\}$), corresponding to a “strongly unbalanced”, “moderately unbalanced”, and “balanced” design, respectively. We compute the power for a number of covariance matrices Σ_a given as follows: For $a = 1, 5, 9$ define

$$\Sigma_a = 10^{-1} \text{diag} \left(\frac{a}{n_1}, \dots, \frac{a}{n_1}, \frac{10-a}{n-n_1}, \dots, \frac{10-a}{n-n_1} \right) \in \mathfrak{C}_{Het},$$

where the first n_1 (and last $n - n_1$, respectively) diagonal entries of each Σ_a are constant. That is, we look at power functions evaluated at covariance matrices under which the subjects in the same group actually have the same variances. [For brevity we do not report power functions for covariance matrices not sharing this property.] For the balanced design, we note that Σ_1 and Σ_9 lead to the same power of each test (but we report all results for completeness), and that Σ_5 corresponds to homoskedasticity.

The critical values are chosen in each case as the smallest critical value guaranteeing size control over \mathfrak{C}_{Het} (implying, of course, that the corresponding tests can have null rejection probabilities smaller than α for the covariance matrices Σ_a considered). The existence of said critical values follows from our theory and is discussed in detail in Example 4.4 for the test statistics UC and HC0-HC4; in particular, all assumptions of Theorems 4.1 are satisfied. For UCR the existence is guaranteed by Part (a) of Theorem 5.1. With regard to the test statistics HC0R-HC4R, note that Assumption 2 is satisfied since $e_i(n) \notin \mathfrak{M}_0^{lin} = \mathfrak{M}_0 = \text{span}((1, \dots, 1)')$ for every $i = 1, \dots, n$ as $n = 30 > k = 2$. This also shows that the sufficient condition for size control (17) is satisfied as $\tilde{\mathfrak{B}} = \mathfrak{M}_0$ is easily verified and since one may set $\mu_0 = 0$. The non-constancy assumption on the test statistics HC0R-HC4R in Theorem 5.1 we have verified numerically. As a consequence, all assumptions of Theorem 5.1 are satisfied.

We note that some of the test statistics differ from each other only by a known multiplicative constant and hence are equivalent in the sense that they give rise to the same test *when the respective smallest size-controlling critical value is employed*, see Remarks 4.3 and 5.2: In the unbalanced case ($n_1 \in \{3, 9\}$), HC0 and HC1 are equivalent in this sense, as are HC0R-HC4R (the latter is so since $\tilde{h}_{ii} = 1/n$ which does not depend on i). In the balanced case ($n = 15$), UC and HC0-HC4 are all equivalent, and the same is true for UCR and HC0R-HC4R as is not difficult to see. Furthermore, in the balanced as well as in the unbalanced case, the rejection regions of the tests based on UC and UCR coincide essentially (i.e., up to a $\lambda_{\mathbb{R}^n}$ -null set) as a consequence of the relationship established in Section 5.1.1. In particular, it follows that in the balanced case, *all* tests considered (essentially) coincide. We nevertheless compute the power functions for each of the tests separately without making use of the noted equivalencies; this provides a double-check of our numerical results.⁴⁶

Numerically the critical values were determined through the implementation of Algorithms 1 and 3 in the R-package **hrt** (Preinerstorfer (2021)) version 1.0.0, and the power functions were computed with the implementation of the algorithm by Davies (1980) in the R-package **CompQuadForm** (Duchesne and de Micheaux (2010)) version 1.4.3; see Appendices E.2 and F.2 for more details. For the sake of illustration, we also report the critical values obtained for every test considered and every balancedness condition in Table 4.

In relation to Table 4 we note that the equivalences discussed before predict, e.g., that the ratio between the entries in the column labeled HC0 and the corresponding entries in the column labeled HC1 should be equal to $n/(n-2) = 30/28 \approx 1.0714$. The ratios computed from the table

⁴⁶The equivalencies mentioned in this paragraph for the two-group-comparison problem hold for general n , n_1 , and n_2 as is easily seen.

n_0	UC	HC0	HC1	HC2	HC3	HC4
3	225.97	26.69	25.63	17.48	11.86	5.43
9	12.70	5.80	5.39	5.10	4.55	4.70
15	4.59	4.91	4.58	4.59	4.28	4.58
n_0	UCR	HC0R	HC1R	HC2R	HC3R	HC4R
3	25.82	3.25	3.14	3.14	3.02	3.13
9	9.05	4.28	4.15	4.14	4.06	4.19
15	4.08	4.23	3.92	4.08	3.95	4.09

Table 4: The smallest size-controlling critical values for comparing the means of two heteroskedastic groups.

are 1.0414, 1.0761, and 1.0721 (for $n_1 = 3, 9, 15$), which is in pretty good agreement (especially if one converts the critical values shown in the table to critical values for the corresponding “ t -test” versions by computing their square roots). The agreement between theoretical and observed ratios for the HC0R-HC4R columns is similar. In the balanced case one can also use the additional equivalences mentioned before and one again finds very good agreement. Similarly, the critical values for UC and UCR in Table 4 are in excellent agreement with their theoretical relationship found in Section 5.1.1. The reason for the small discrepancies observed lies in the fact that the algorithm underlying the computations for Table 4 makes use of a random search algorithm. Concerning Table 4, we also mention that, in the example considered here and for the test statistic HC2, Ibragimov and Müller (2016) prove in their Theorem 1 (see also the discussion preceding that theorem) that the smallest size-controlling critical values are given by 18.51 ($n_1 = 3$), 5.32 ($n_1 = 9$), and 4.60 ($n_1 = 15$), respectively. The numerically determined critical values in Table 4 are reasonably close to these values (after conversion of the critical values to corresponding “ t -test” critical values the maximal difference is about 0.1). Of course, the accuracy of our algorithm could be increased by using more stringent accuracy parameters in the optimization routines underlying the computation of the critical value, but this would come with a longer runtime.

From Table 4 it is clear that for the tests based on *unrestricted* residuals the smallest size-controlling critical values obtained are always larger, sometimes considerably, than $C_{\chi^2, 0.05} \approx 3.8415$, again showing that the latter critical value is not effecting size control. For the tests based on *restricted* residuals the smallest size-controlling critical values sometimes fall below $C_{\chi^2, 0.05}$ in the strongly unbalanced case (which is not completely surprising in view of Section 5.1.2); while in this case $C_{\chi^2, 0.05}$ effects size-control, using the smaller size-controlling critical values given in Table 4 can only be advantageous in terms of power.

That being said, we emphasize a trivial, but important point, namely that comparing the magnitudes of size-controlling critical values relating to *different* test statistics is not very meaningful and, in particular, *not* a valid way of comparing the quality of the resulting tests. That is, while it may be tempting to infer from Table 4 that the HC0 test should be considerably more conservative than the HC4 test, or that the UC test should be considerably more conservative

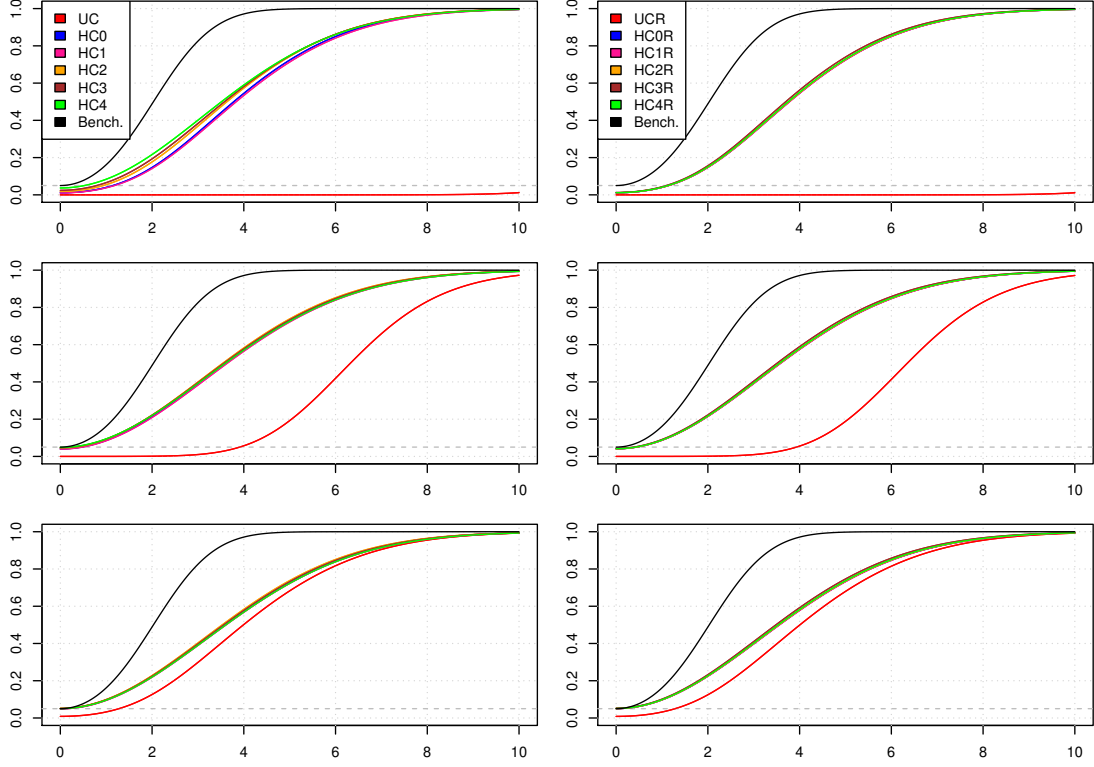


Figure 1: Power functions for $n_1 = 3$. Left column: tests based on unrestricted residuals (cf. legend). Right column: tests based on restricted residuals (cf. legend). The rows corresponds to Σ_a for $a = 1, 5, 9$ from top to bottom. The abscissa shows δ .

than the UCR test, such a conclusion would be false and not warranted at all (in particular, recall that UC and UCR in fact result in (essentially) the same test if the critical values from Table 4 are being used). While this would be correct if the critical values were all meant to be used with the same test statistic (which they are not), critical values belonging to different test statistics can certainly not be compared in such a way. Instead, one has to compare the corresponding power functions, which is what we shall do next.

The power functions are shown in Figure 1 (“strongly unbalanced”, $n_1 = 3$), Figure 2 (“moderately unbalanced”, $n_1 = 9$), and Figure 4 (“balanced”, $n_1 = 15$), where only the first two figures are shown in the main text, and the last figure (in which all power functions lie “on top of each other”) is available in Appendix F.3.

The power functions illustrate that the testing problem is getting easier, (i.e., power gets closer to the oracle benchmark), for more balanced design, which has intuitive appeal. Except for the strongly unbalanced case ($n_1 = 3$), the power loss of the tests based on HC0-HC4 and HC0R-HC4R relative to the oracle benchmark is surprisingly small. In the unbalanced cases ($n_1 \in \{3, 9\}$) the HC0-HC4-based tests behave all very similarly, with the power functions of

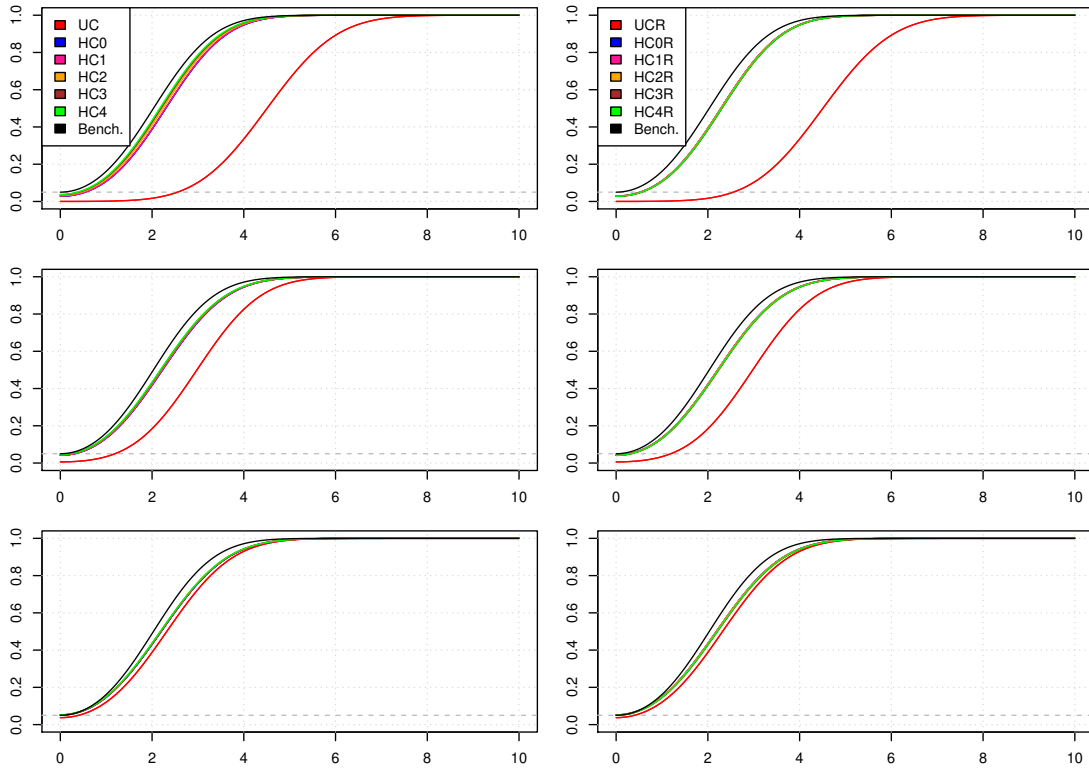


Figure 2: Power functions for $n_1 = 9$. Left column: tests based on unrestricted residuals (cf. legend). Right column: tests based on restricted residuals (cf. legend). The rows corresponds to Σ_a for $a = 1, 5, 9$ from top to bottom. The abscissa shows δ .

the HC0- and HC1-based test being virtually indistinguishable (as they should in view of the before discussed equivalence). The UC-based test shows markedly worse power performance. Similarly, the HC0R-HC4R-based tests have virtually indistinguishable power functions (as they should because of the before discussed equivalence). The UCR-based test again is inferior (and its power function coincides with the one of UC as mentioned before). There appears also to be little difference between basing the test statistics on unrestricted or restricted residuals in this example. In the balanced case we know that *all* the tests have exactly the same power function in view of our earlier discussion. This is visible in Figure 4 in Appendix F.3. Also the different forms of heteroskedasticity considered seem not to have much effect on the power functions (when expressed as a function of δ), except for UC and UCR in the unbalanced cases.

Hence, within the scenario considered in this section, perhaps the most important conclusion concerning the choice of a test statistic appears to be to avoid UC and UCR. Everything apart from that, i.e., whether one uses unrestricted or restricted residuals to construct the test or which specific heteroskedasticity-correction one decides to use, seems to be a comparably irrelevant part of the problem once the right (i.e., smallest size-controlling) critical value is used. We shall see in the next subsection that this conclusion very much depends on the scenario considered here and does not generalize beyond, illustrating the danger of drawing conclusions from a limited numerical study.

9.2.2 A high-leverage design matrix

In this section, we consider testing $\beta_2 = 0$ in a model with intercept and a single regressor $x = (10, \cos(2), \cos(3), \dots, \cos(n))'$. Obviously, the regressor has a dominant first coordinate, leading to diagonal elements h_{ii} of $X(X'X)^{-1}X'$ such that the ratio of largest to smallest h_{ii} is roughly 26. Hence, the design matrix X provides (on purpose) an extreme case, which leads to quite interesting results. We consider again the case $n = 30$ and $\alpha = 0.05$, but now show power functions for Σ_a^* , $a = 0, \dots, 4$, where

$$\Sigma_a^* = n^{-1} \text{diag} \left(7a + 1, \frac{n - 7a - 1}{n - 1}, \dots, \frac{n - 7a - 1}{n - 1} \right) \in \mathfrak{C}_{Het}.$$

Note that $\Sigma_0^* = n^{-1}I_n$ and that increasing a from 0 to 4 leads to covariance matrices that approach the degenerate matrix $e_1(n)e_1(n)'$. All conditions in Theorems 4.1 and 5.1 are seen to be satisfied in this example: As no vector $e_i(n)$ belongs to $\text{span}(X)$ (and thus also not to \mathfrak{M}_0^{lin}), Assumptions 1 and 2 as well as the sufficient condition for size control (11) are obviously satisfied. The size control conditions (13) and (17) have been checked numerically, as has been the condition that none of the test statistics HC0R-HC4R is constant on $\mathbb{R}^n \setminus \tilde{\mathcal{B}}$.

As in the preceding subsection, the critical values for each test statistic are again chosen as the smallest critical value guaranteeing size control over \mathfrak{C}_{Het} and they are presented in Table 5 below. [Existence follows from our theory since all assumptions are satisfied as noted before.] For their computation the same algorithms were used as in Section 9.2.1, with a similar statement

applying to the numerical routines used for computing the power functions. Note that the critical values for the test statistics UC, HC0-HC3 are large, reflecting the high-leverage in the design matrix; an exception is HC4, the reason being that some of the HC4-weights are considerably larger than the weights for HC0-HC3. Similarly as in the preceding subsection, the tests based on HC0 and HC1 coincide, and the same is true for the tests based on HC0R-HC4R, see Remarks 4.3 and 5.2. It is easily checked that the ratios of the respective critical values provided in Table 5 are in good agreement with the theoretical ratios predicted by theory. Similarly, the tests based on UC and UCR coincide, and the critical values for UC and UCR in Table 5 are in excellent agreement with their theoretical relationship found in Section 5.1.1.

Table 5 shows that in this example the smallest size-controlling critical values are – except in one case – always larger, sometimes considerably larger, than $C_{\chi^2,0.05} \approx 3.8415$, once more showing that the latter critical value is not effecting size control in general. In the exceptional case, namely when the HC4 test statistic is used, $C_{\chi^2,0.05}$ is considerably larger than the smallest size-controlling critical value, which is 1.12; while in this case $C_{\chi^2,0.05}$ effects size-control, using the smaller size-controlling critical value 1.12 can only be advantageous in terms of power.

UC	HC0	HC1	HC2	HC3	HC4
217.58	355.56	333.31	121.89	29.34	1.12
UCR	HC0R	HC1R	HC2R	HC3R	HC4R
25.69	5.41	5.45	5.34	5.29	5.44

Table 5: Smallest size-controlling critical values for the high-leverage design matrix.

The power functions, when the size-controlling critical values from Table 5 are being used, are shown in Figure 3. Again, as predicted by theory, the power functions of the tests based on HC0 and HC1 shown in Figure 3 coincide, as do the power functions of the tests based on HC0R-HC4R; the same is true for the power functions of the tests based on UC and UCR. The figure furthermore shows that in the setting considered here, there is now a marked difference between tests based on HC0-HC4 and on HC0R-HC4R, respectively: the power of the tests based on HC0R-HC4R is nowhere greater than α , their power function being even non-monotonic, whereas the tests based on HC0-HC4 have increasing power as a function of δ . In contrast to the example considered in the preceding subsection, the power functions of the tests based on HC0-HC4 and on UC are now all markedly different and typically intersect, an exception being the case of Σ_4^* where the test based on UC offers the highest power for that covariance matrix. Overall, however, there is no clear ranking between the tests using unrestricted residuals in the example considered here, although we note that the test based on UC (or, equivalently, on UCR) performs very badly in the case of Σ_0^* . This is not surprising as Σ_0^* corresponds to homoskedasticity and the critical value used here is much larger than the classical critical value one would use given knowledge of this homoskedasticity. Furthermore, and in contrast to the results in the preceding subsection, the different forms of heteroskedasticity considered have a noticeable effect on the power functions. The main takeaway is that tests based on HC0R-HC4R

(and probably on UC and UCR) should rather be avoided.

10 Conclusion

The usual heteroskedasticity robust tests statistics such as T_{Het} (using HC0-HC4 weights) or \tilde{T}_{Het} (using HC0R-HC4R weights), used in conjunction with conventional critical values obtained from the asymptotic null distribution, are often plagued by overrejection under the null. This has been clearly documented in the literature for T_{Het} , and is shown numerically for \tilde{T}_{Het} (as well as for T_{Het}) in Section 9 above. Not surprisingly, similar observations apply to the “uncorrected” test statistics T_{uc} and \tilde{T}_{uc} . We show theoretically that all these test statistics can be size-controlled under quite weak conditions by an appropriate choice of critical values.

From the above discussion and the numerical results in Section 9 it transpires that smallest size-controlling critical values rather than conventional critical values should be used in order to avoid the risk of overrejection. For the computation of smallest size-controlling critical values we provide algorithms which have been implemented in the R-package **hrt** (Preinerstorfer (2021)) and thus are readily available for the user.

An additional advantage from using smallest size-controlling critical values over conventional critical values is that this typically leads to improved power in instances, where conventional critical values lead to underrejection (i.e., lead to worst-case rejection probability under the null less than the nominal significance level) as is sometimes the case; see Sections 5.1.2 and 9.2.

If smallest size-controlling critical values are adopted (as they should), the numerical results in Section 9 suggest that the test statistic \tilde{T}_{Het} (with the usual weights HC0R-HC4R) should be avoided, as the resulting tests may have very poor power properties (see the example in Section 9.2.2). The test statistic T_{Het} seems to perform better, with no clear ranking emerging with regards to the weights HC0-HC4 being used. The “uncorrected” test statistics T_{uc} and \tilde{T}_{uc} appear to be inferior to T_{Het} in almost all of the numerical examples considered. We also point out that – when using smallest size-controlling critical values – the tests based on T_{Het} employing the HC0 and the HC1 weights, respectively, in fact coincide; and the same holds for tests based on \tilde{T}_{Het} employing the HC0R and the HC1R weights, respectively. Also the tests based on T_{uc} and \tilde{T}_{uc} then (essentially) coincide. See Remarks 4.3, 5.2, and Section 5.1.1 as well as the pertaining discussion in Section 9 for more information, including additional equivalencies when the design matrix X and the restriction R have certain special properties.

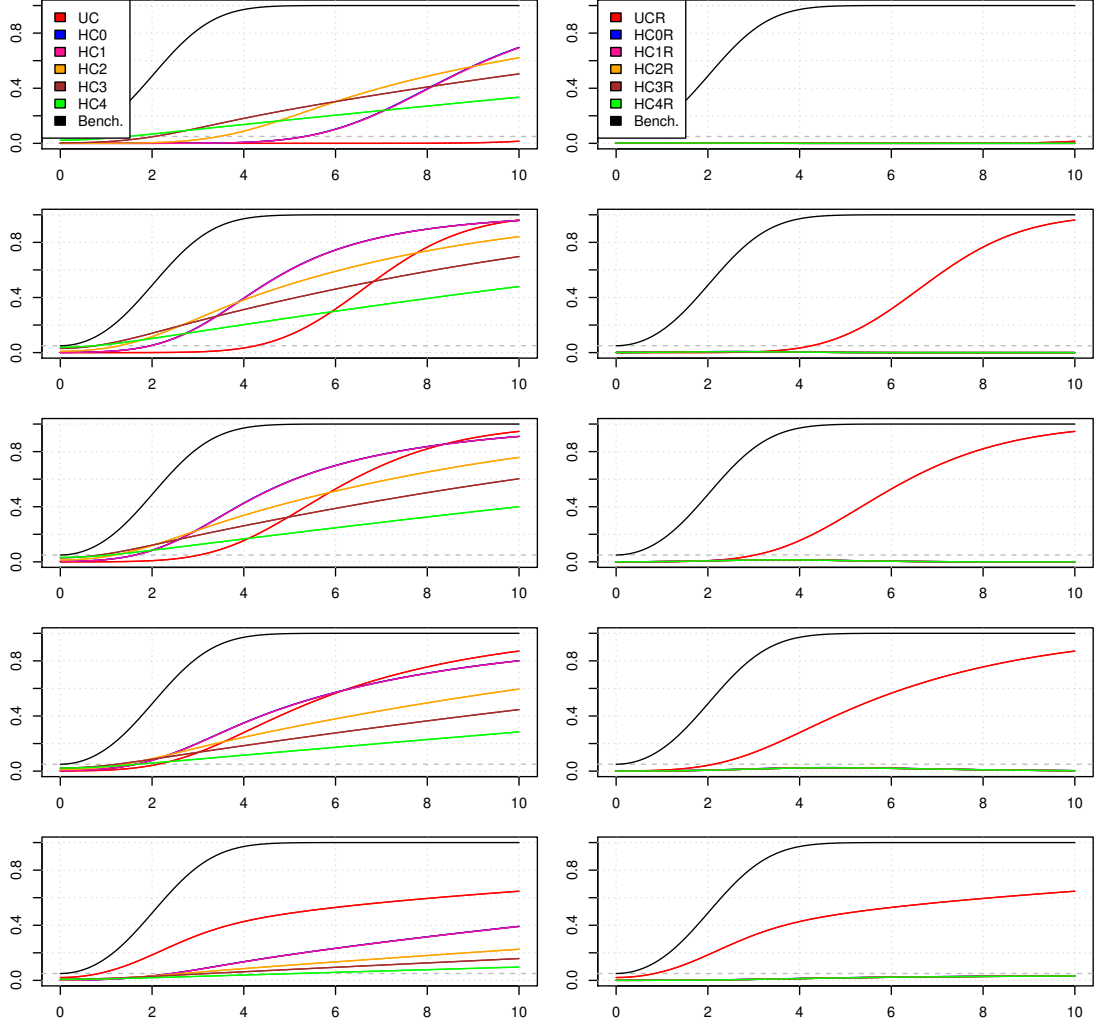


Figure 3: Power functions for the design matrix considered in Section 9.2.2. Left column: tests based on unrestricted residuals (cf. legend). Right column: tests based on restricted residuals (cf. legend). The rows from top to bottom correspond to Σ_a^* for $a = 0, 1, 2, 3, 4$, the case $a = 0$ corresponding to homoskedasticity. The abscissa shows δ .

A Appendix: Size control over other heteroskedasticity models

As already noted earlier, if size control is possible over \mathfrak{C}_{Het} , then the same is true over *any* conceivable class of heteroskedasticity structures, since these can (possible after normalization) be cast as a subset \mathfrak{C} of \mathfrak{C}_{Het} ; and, in fact, any critical value delivering size control over \mathfrak{C}_{Het} also delivers size control over any such \mathfrak{C} , but even smaller critical values may already suffice for size control over \mathfrak{C} . Also, for some heteroskedasticity models $\mathfrak{C} \subseteq \mathfrak{C}_{Het}$, the sufficient conditions employed in Theorems 4.1 and 5.1 (which imply size control over \mathfrak{C}_{Het}) may be unnecessarily restrictive, if one wants to establish size control over \mathfrak{C} only. For this reason, we show in the following how the general theory laid out in Section 5 of Pötscher and Preinerstorfer (2018) can be used to derive size control results tailored to various subsets \mathfrak{C} by exemplarily treating the cases $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$ and $\mathfrak{C} = \mathfrak{C}_{Het, \tau_*}$ defined below. Size control results over other choices of \mathfrak{C} can be derived from the results in Section 5 of Pötscher and Preinerstorfer (2018) in a similar manner, see Subsection A.1.2 further below for some discussion. Here $\mathfrak{C}_{(n_1, \dots, n_m)}$ is defined as follows: Let $m \in \mathbb{N}$, and let $n_j \in \mathbb{N}$ for $j = 1, \dots, m$ satisfy $\sum_{j=1}^m n_j = n$. Set $n_j^+ = \sum_{l=1}^j n_l$ and define

$$\mathfrak{C}_{(n_1, \dots, n_m)} = \left\{ \text{diag}(\tau_1^2, \dots, \tau_n^2) \in \mathfrak{C}_{Het} : \tau_{n_{j-1}^+ + 1}^2 = \dots = \tau_{n_j^+}^2 \text{ for } j = 1, \dots, m \right\}$$

with the convention that $n_0^+ = 0$. This may be a natural heteroskedasticity model when the observations come from m groups and when it is reasonable to assume homoskedasticity within groups.⁴⁷ Note that in case $n_j = 1$ for all j , then $m = n$ and $\mathfrak{C}_{(n_1, \dots, n_m)} = \mathfrak{C}_{Het}$ hold; and in case $m = 1$ we have $\mathfrak{C}_{(n_1, \dots, n_m)} = \{n^{-1}I_n\}$, i.e., we have homoskedasticity. Furthermore, $\mathfrak{C}_{Het, \tau_*}$ is given by

$$\mathfrak{C}_{Het, \tau_*} = \left\{ \text{diag}(\tau_1^2, \dots, \tau_n^2) \in \mathfrak{C}_{Het} : \tau_i^2 \geq \tau_*^2 \text{ for all } i \right\},$$

where the lower bound τ_* , $0 < \tau_* < n^{-1/2}$, is set by the user.

A.1 Size control results for T_{Het} and T_{uc}

A.1.1 Size control over $\mathfrak{C}_{(n_1, \dots, n_m)}$

Proofs of the results in this subsection can be found in Appendix C.

Theorem A.1. *Let $m \in \mathbb{N}$, and let $n_j \in \mathbb{N}$ for $j = 1, \dots, m$ satisfy $\sum_{j=1}^m n_j = n$. Then:*

(a) *For every $0 < \alpha < 1$ there exists a real number $C(\alpha)$ such that*

$$\sup_{\mu_0 \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}_{(n_1, \dots, n_m)}} P_{\mu_0, \sigma^2 \Sigma}(T_{uc} \geq C(\alpha)) \leq \alpha \quad (21)$$

⁴⁷As long as we assume that the grouping is known, there is little loss of generality to assume that the elements belonging to the same group are numbered contiguously, since we otherwise only need to relabel the data.

holds, provided that

$$\text{span}(\{e_i(n) : i \in (n_{j-1}^+, n_j^+]\}) \not\subseteq \text{span}(X) \text{ for every } j = 1, \dots, m \text{ with } (n_{j-1}^+, n_j^+] \cap I_1(\mathfrak{M}_0^{lin}) \neq \emptyset. \quad (22)$$

Furthermore, under condition (22), even equality can be achieved in (21) by a proper choice of $C(\alpha)$, provided $\alpha \in (0, \alpha^*] \cap (0, 1)$ holds, where $\alpha^* = \sup_{C \in (C^*, \infty)} \sup_{\Sigma \in \mathfrak{C}_{(n_1, \dots, n_m)}} P_{\mu_0, \Sigma}(T_{uc} \geq C)$ is positive and where C^* is defined as in Lemma 5.11 of Pötscher and Preinerstorfer (2018) with $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$, $T = T_{uc}$, $N^\dagger = \text{span}(X)$, and $\mathcal{L} = \mathfrak{M}_0^{lin}$ (with neither α^* nor C^* depending on the choice of $\mu_0 \in \mathfrak{M}_0$).

(b) Suppose Assumption 1 is satisfied. Then for every $0 < \alpha < 1$ there exists a real number $C(\alpha)$ such that

$$\sup_{\mu_0 \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}_{(n_1, \dots, n_m)}} P_{\mu_0, \sigma^2 \Sigma}(T_{Het} \geq C(\alpha)) \leq \alpha \quad (23)$$

holds, provided that

$$\text{span}(\{e_i(n) : i \in (n_{j-1}^+, n_j^+]\}) \not\subseteq B \text{ for every } j = 1, \dots, m \text{ with } (n_{j-1}^+, n_j^+] \cap I_1(\mathfrak{M}_0^{lin}) \neq \emptyset. \quad (24)$$

Furthermore, under condition (24), even equality can be achieved in (23) by a proper choice of $C(\alpha)$, provided $\alpha \in (0, \alpha^*] \cap (0, 1)$ holds, where $\alpha^* = \sup_{C \in (C^*, \infty)} \sup_{\Sigma \in \mathfrak{C}_{(n_1, \dots, n_m)}} P_{\mu_0, \Sigma}(T_{Het} \geq C)$ is positive and where C^* is defined as in Lemma 5.11 of Pötscher and Preinerstorfer (2018) with $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$, $T = T_{Het}$, $N^\dagger = B$, and $\mathcal{L} = \mathfrak{M}_0^{lin}$ (with neither α^* nor C^* depending on the choice of $\mu_0 \in \mathfrak{M}_0$).

(c) Under the assumptions of Part (a) (Part (b), respectively) implying existence of a critical value $C(\alpha)$ satisfying (21) ((23), respectively), a smallest critical value, denoted by $C_\diamond(\alpha)$, satisfying (21) ((23), respectively) exists for every $0 < \alpha < 1$. And $C_\diamond(\alpha)$ corresponding to Part (a) (Part (b), respectively) is also the smallest among the critical values leading to equality in (21) ((23), respectively) whenever such a critical value exists. [Although $C_\diamond(\alpha)$ corresponding to Part (a) and (b), respectively, will typically be different, we use the same symbol.]

It is easy to see that the discussion in the first paragraph following Theorem 4.1 applies mutatis mutandis also to the above theorem. Similarly, Remarks 4.2, 4.3, 4.4, 4.6, 4.9, and Proposition 4.5 carry over. Furthermore, we have the following result corresponding to Proposition 4.7:

Proposition A.2. (a) If (22) is violated, then $\sup_{\Sigma \in \mathfrak{C}_{(n_1, \dots, n_m)}} P_{\mu_0, \sigma^2 \Sigma}(T_{uc} \geq C) = 1$ for every choice of critical value C , every $\mu_0 \in \mathfrak{M}_0$, and every $\sigma^2 \in (0, \infty)$ (implying that size equals 1 for every C). As a consequence, the sufficient condition for size control (22) in Part (a) of Theorem A.1 is also necessary.

(b) Suppose Assumption 1 is satisfied.⁴⁸ If (22) is violated, then $\sup_{\Sigma \in \mathfrak{C}_{(n_1, \dots, n_m)}} P_{\mu_0, \sigma^2 \Sigma}(T_{Het} \geq C) = 1$ for every choice of critical value C , every $\mu_0 \in \mathfrak{M}_0$, and every $\sigma^2 \in (0, \infty)$ (implying

⁴⁸If this assumption is violated then T_{Het} is identically zero, an uninteresting trivial case.

that size equals 1 for every C). [In case X and R are such that $B = \text{span}(X)$, conditions (22) and (24) coincide; hence the sufficient condition for size control (24) in Part (b) of Theorem A.1 is then also necessary in this case.]

Remark A.3. (*Homoskedasticity*) Theorem A.1 allows also for the case $m = 1$, in which case $\mathfrak{C}_{(n_1, \dots, n_m)} = \{n^{-1}I_n\}$, i.e., errors are homoskedastic. In this case it is easy to see that the sufficient conditions for size control in the theorem are trivially satisfied and size control for T_{Het} (and T_{uc}) is possible.⁴⁹ Of course, this is in line with the fact that T_{Het} and T_{uc} are obviously pivotal under the null if the errors are homoskedastic.

Remark A.4. (*Behrens-Fisher problem*) Consider again the problem of testing the equality of the means of two independent normal populations as in Example 4.4 with the only difference that the variance within each of the two groups is now assumed to be constant, i.e., the heteroskedasticity model used is now given by $\mathfrak{C}_{(n_1, n_2)}$, where $n_1 \geq 2$ and $n_2 \geq 2$ are the group sizes. This is the celebrated Behrens-Fisher problem. The square of the two-sample t-statistic t_{FB} , say, often used in this context coincides with T_{Het} for the choice $d_i = (1 - h_{ii})^{-1}$. The size controllability of T_{Het} over \mathfrak{C}_{Het} established in Example 4.4 therefore a fortiori implies size controllability of T_{Het} (and hence of t_{FB}^2) over $\mathfrak{C}_{(n_1, n_2)}$. Of course, this does not add anything new, since it is known that under the null hypothesis $|t_{FB}|$ is stochastically not larger than a t -distributed random variable with $\min(n_1, n_2) - 1$ degrees of freedom, see Mickey and Brown (1966). For more on the Behrens-Fisher problem see Kim and Cohen (1998), Ruben (2002), Lehmann and Romano (2005), Belloni and Didier (2008), and the references cited therein.

A.1.2 Further size control results

In this subsection it is understood that Assumption 1 is maintained when discussing results relating to T_{Het} .

(i) Given a heteroskedasticity model \mathfrak{C} (i.e., $\emptyset \neq \mathfrak{C} \subseteq \mathfrak{C}_{Het}$), with the property that $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C})$ is empty (where the collection $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C})$ is defined on p. 421 of Pötscher and Preinerstorfer (2018), see also Appendix B further below), the tests based on T_{uc} and T_{Het} are always size controllable over \mathfrak{C} . This follows from Corollary 5.6 and Remark 5.7 in Pötscher and Preinerstorfer (2018). In fact, exact size control is then possible for every $\alpha \in (0, 1)$ as a consequence of Proposition 5.12 in the same reference upon noting that then $C^* = -\infty$ and $\alpha^* = 1$ hold. [We note in passing that for such a heteroskedasticity model \mathfrak{C} the size of the rejection region $\{T_{uc} \geq C\}$ ($\{T_{Het} \geq C\}$, respectively) is less than 1 for every $C > 0$ (this follows from Proposition 5.2 and Remark 5.4 in Pötscher and Preinerstorfer (2018) as well as Part 6 of Lemma 5.15 in Preinerstorfer and Pötscher (2016)).]⁵⁰

⁴⁹A related but slightly different argument proceeds by directly noting from its definition that $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$ is empty in case $m = 1$ (cf. Appendix B), and then to appeal to Remark 5.7 (or Proposition 5.12) in Pötscher and Preinerstorfer (2018).

⁵⁰The verification of the assumptions in Corollary 5.6 and Propositions 5.2 and 5.12 of Pötscher and Preinerstorfer (2018) proceeds as in the proofs of Theorems 4.1 and A.1.

(ii) A particular instance of the situation described in (i) is provided by heteroskedasticity models \mathfrak{C} that are subsets of a set of the form $\mathfrak{C}_{Het, \tau_*}$ ($0 < \tau_* < n^{-1/2}$), as in this case $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C})$ is easily seen to be empty.

(iii) More generally, the tests based on T_{uc} (on T_{Het} , respectively) are size controllable over a heteroskedasticity model \mathfrak{C} , provided any $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C})$ is not contained in $\text{span}(X)$ (\mathfrak{B} , respectively). This follows easily from Corollary 5.6 and Proposition 5.12 in Pötscher and Preinerstorfer (2018), the latter proposition also providing an exact size result, which we refrain from spelling out in detail. Again there is a (partial) converse: If an $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C})$ exists with $\mathcal{S} \subseteq \text{span}(X)$, then the size over \mathfrak{C} of the rejection region $\{T_{uc} \geq C\}$ ($\{T_{Het} \geq C\}$, respectively) is equal to 1; see Theorem 3.1 in Pötscher and Preinerstorfer (2019). Furthermore, lower bounds for critical values that lead to size less than 1 (in particular, for size-controlling critical values) can be had with the help of Corollary 5.17 in Preinerstorfer and Pötscher (2016), Lemma 5.11 and Proposition 5.12 in Pötscher and Preinerstorfer (2018), or Lemma 4.1 in Pötscher and Preinerstorfer (2019).

A.2 Size control results for \tilde{T}_{Het} and \tilde{T}_{uc}

The proof of the subsequent theorem is given in Appendix D. We note that the first statement in Part (a) of the subsequent theorem is actually trivial, since \tilde{T}_{uc} is bounded as has been shown in Section 5.1.1.

Theorem A.5. *Let $m \in \mathbb{N}$, and let $n_j \in \mathbb{N}$ for $j = 1, \dots, m$ satisfy $\sum_{j=1}^m n_j = n$. Then:*

(a) *For every $0 < \alpha < 1$ there exists a real number $C(\alpha)$ such that*

$$\sup_{\mu_0 \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}_{(n_1, \dots, n_m)}} P_{\mu_0, \sigma^2 \Sigma}(\tilde{T}_{uc} \geq C(\alpha)) \leq \alpha \quad (25)$$

holds. Furthermore, even equality can be achieved in (25) by a proper choice of $C(\alpha)$, provided $\alpha \in (0, \alpha^] \cap (0, 1)$ holds, where $\alpha^* = \sup_{C \in (C^*, \infty)} \sup_{\Sigma \in \mathfrak{C}_{(n_1, \dots, n_m)}} P_{\mu_0, \Sigma}(\tilde{T}_{uc} \geq C)$ and where C^* is defined as in Lemma 5.11 of Pötscher and Preinerstorfer (2018) with $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$, $T = \tilde{T}_{uc}$, $N^\dagger = \mathfrak{M}_0$, and $\mathcal{L} = \mathfrak{M}_0^{lin}$ (with neither α^* nor C^* depending on the choice of $\mu_0 \in \mathfrak{M}_0$).*

(b) *Suppose Assumption 2 is satisfied. Suppose further that \tilde{T}_{Het} is not constant on $\mathbb{R}^n \setminus \tilde{\mathfrak{B}}$.⁵¹ Then for every $0 < \alpha < 1$ there exists a real number $C(\alpha)$ such that*

$$\sup_{\mu_0 \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}_{(n_1, \dots, n_m)}} P_{\mu_0, \sigma^2 \Sigma}(\tilde{T}_{Het} \geq C(\alpha)) \leq \alpha \quad (26)$$

holds, provided that for some $\mu_0 \in \mathfrak{M}_0$ (and hence for all $\mu_0 \in \mathfrak{M}_0$)

$$\mu_0 + \text{span}(\{e_i(n) : i \in (n_{j-1}^+, n_j^+]\}) \not\subseteq \tilde{\mathfrak{B}} \quad \text{for every } j = 1, \dots, m \text{ with } (n_{j-1}^+, n_j^+] \cap I_1(\mathfrak{M}_0^{lin}) \neq \emptyset. \quad (27)$$

⁵¹Cf. Footnote 28.

Furthermore, under condition (27), even equality can be achieved in (26) by a proper choice of $C(\alpha)$, provided $\alpha \in (0, \alpha^*] \cap (0, 1)$ holds, where $\alpha^* = \sup_{C \in (C^*, \infty)} \sup_{\Sigma \in \mathfrak{C}_{(n_1, \dots, n_m)}} P_{\mu_0, \Sigma}(\tilde{T}_{Het} \geq C)$ and where C^* is defined as in Lemma 5.11 of Pötscher and Preinerstorfer (2018) with $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$, $T = \tilde{T}_{Het}$, $N^\dagger = \tilde{\mathbf{B}}$, and $\mathcal{L} = \mathfrak{M}_0^{lin}$ (with neither α^* nor C^* depending on the choice of $\mu_0 \in \mathfrak{M}_0$).

(c) Under the assumptions of Part (a) (Part (b), respectively) implying existence of a critical value $C(\alpha)$ satisfying (25) ((26), respectively), a smallest critical value, denoted by $C_\diamond(\alpha)$, satisfying (25) ((26), respectively) exists for every $0 < \alpha < 1$.⁵² And $C_\diamond(\alpha)$ corresponding to Part (a) (Part (b), respectively) is also the smallest among the critical values leading to equality in (25) ((26), respectively) whenever such a critical value exists. [Although $C_\diamond(\alpha)$ corresponding to Part (a) and (b), respectively, will typically be different, we use the same symbol.]

It is easy to see that the discussion in the first paragraph following Theorem 5.1 applies mutatis mutandis also to the above theorem. Similarly, Remarks 5.2, 5.3, 5.5, 5.7, and Proposition 5.4 carry over.

A discussion of size control results for \tilde{T}_{uc} and \tilde{T}_{Het} over other choices of \mathfrak{C} based on the results in Section 5 of Pötscher and Preinerstorfer (2018) can also be given (cf. the discussion in Subsection A.1.2), but we refrain from spelling out the details. We only note that the test based on \tilde{T}_{Het} is always size controllable over $\mathfrak{C}_{Het, \tau_*}$ ($0 < \tau_* < n^{-1/2}$), and the same is trivially true for \tilde{T}_{uc} .

B Appendix: Characterization of $\mathbb{J}(\mathcal{L}, \mathfrak{C})$ for $\mathfrak{C} = \mathfrak{C}_{Het}$ and $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$

A key ingredient in the proof of size control results such as Theorem 4.1 or 5.1 is a certain collection $\mathbb{J}(\mathcal{L}, \mathfrak{C})$ of linear subspaces of \mathbb{R}^n introduced in Pötscher and Preinerstorfer (2018). For the convenience of the reader we reproduce this definition, specialized to the present setting, below. The leading case in the applications will be the case $\mathcal{L} = \mathfrak{M}_0^{lin}$.

Definition B.1. Let \mathfrak{C} be a heteroskedasticity model, i.e., $\mathfrak{C} \subseteq \mathfrak{C}_{Het}$. Given a linear subspace \mathcal{L} of \mathbb{R}^n with $\dim(\mathcal{L}) < n$ and an element $\Sigma \in \mathfrak{C}$, we let

$$\mathcal{L}(\Sigma) = \frac{\Pi_{\mathcal{L}^\perp} \Sigma \Pi_{\mathcal{L}^\perp}}{\|\Pi_{\mathcal{L}^\perp} \Sigma \Pi_{\mathcal{L}^\perp}\|}$$

and $\mathcal{L}(\mathfrak{C}) = \{\mathcal{L}(\Sigma) : \Sigma \in \mathfrak{C}\}$. Furthermore, we define

$$\mathbb{J}(\mathcal{L}, \mathfrak{C}) = \{\text{span}(\bar{\Sigma}) : \bar{\Sigma} \in \text{cl}(\mathcal{L}(\mathfrak{C})), \text{rank}(\bar{\Sigma}) < n - \dim(\mathcal{L})\},$$

where the closure $\text{cl}(\cdot)$ is to be understood w.r.t. $\mathbb{R}^{n \times n}$.

⁵²Note that there are in fact no assumptions for Part (a). We have chosen this formulations for reasons of brevity.

Recalling the definition of $I_0(\mathcal{L})$, it is easy to see that $I_0(\mathcal{L}) = \{i : 1 \leq i \leq n, \pi_{\mathcal{L}^\perp, i} = 0\}$ holds, where $\pi_{\mathcal{L}^\perp, i}$ denotes the i -th column of $\Pi_{\mathcal{L}^\perp}$. Also recall that $I_1(\mathcal{L})$ is nonempty in case $\dim(\mathcal{L}) < n$ holds. The characterization of $\mathbb{J}(\mathcal{L}, \mathfrak{C}_{Het})$ is now given in the next proposition.

Proposition B.1. *Suppose $\dim(\mathcal{L}) < n$ holds. Then the set $\mathbb{J}(\mathcal{L}, \mathfrak{C}_{Het})$ is given by*

$$\{\text{span}(\{\pi_{\mathcal{L}^\perp, i} : i \in I\}) : \emptyset \neq I \subseteq I_1(\mathcal{L}), \dim(\text{span}(\{\pi_{\mathcal{L}^\perp, i} : i \in I\})) < n - \dim(\mathcal{L})\}. \quad (28)$$

This proposition is a special case of Proposition B.2 given below since \mathfrak{C}_{Het} coincides with $\mathfrak{C}_{(n_1, \dots, n_m)}$ in case $m = n$ and $n_j = 1$ for all $j = 1, \dots, m$.

We next turn to the characterization of $\mathbb{J}(\mathcal{L}, \mathfrak{C}_{(n_1, \dots, n_m)})$, where $\mathfrak{C}_{(n_1, \dots, n_m)}$ has been defined in Appendix A. Here $m \in \mathbb{N}$, and $n_j \in \mathbb{N}$ for $j = 1, \dots, m$ satisfy $\sum_{j=1}^m n_j = n$. Consider the partition of the set $\{1, \dots, n\}$ into the intervals $(n_0^+, n_1^+]$, $(n_1^+, n_2^+]$, ..., $(n_{m-1}^+, n_m^+]$ where n_j^+ has been defined in Appendix A. Let $\mathbf{I}_{(n_1, \dots, n_m)}$ consist of all non-empty subsets I of $\{1, \dots, n\}$ that can be represented as a union of intervals of the form $(n_{j-1}^+, n_j^+]$.

Proposition B.2. *Suppose $\dim(\mathcal{L}) < n$ holds. Let $m \in \mathbb{N}$, and let $n_j \in \mathbb{N}$ for $j = 1, \dots, m$ satisfy $\sum_{j=1}^m n_j = n$. Then the set $\mathbb{J}(\mathcal{L}, \mathfrak{C}_{(n_1, \dots, n_m)})$ is given by*

$$\{\text{span}(\{\pi_{\mathcal{L}^\perp, i} : i \in I\}) : I \in \mathbf{I}_{(n_1, \dots, n_m)}, \emptyset \neq I \cap I_1(\mathcal{L}), \dim(\text{span}(\{\pi_{\mathcal{L}^\perp, i} : i \in I\})) < n - \dim(\mathcal{L})\}. \quad (29)$$

Note that in (29) we have $\text{span}(\{\pi_{\mathcal{L}^\perp, i} : i \in I\}) = \text{span}(\{\pi_{\mathcal{L}^\perp, i} : i \in I \cap I_1(\mathcal{L})\})$.

Proof: Suppose \mathcal{S} is an element of $\mathbb{J}(\mathcal{L}, \mathfrak{C}_{(n_1, \dots, n_m)})$. Then there exist a sequence $\Sigma_m \in \mathfrak{C}_{(n_1, \dots, n_m)}$ such that $\Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp} / \|\Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp}\|$ converges to a limit $\bar{\Sigma}$, say, in $\mathbb{R}^{n \times n}$ with $\text{span}(\bar{\Sigma}) = \mathcal{S}$. Now,

$$\begin{aligned} \Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp} / \|\Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp}\| &= \|\Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp}\|^{-1} \sum_{i=1}^n \tau_i^2(m) \pi_{\mathcal{L}^\perp, i} \pi'_{\mathcal{L}^\perp, i} \\ &= \|\Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp}\|^{-1} \sum_{j=1}^m \sum_{i \in (n_{j-1}^+, n_j^+]} \tau_i^2(m) \pi_{\mathcal{L}^\perp, i} \pi'_{\mathcal{L}^\perp, i} \\ &= \sum_{j: (n_{j-1}^+, n_j^+] \cap I_1(\mathcal{L}) \neq \emptyset} \|\Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp}\|^{-1} \tau_{n_j^+}^2(m) \sum_{i \in (n_{j-1}^+, n_j^+]} \pi_{\mathcal{L}^\perp, i} \pi'_{\mathcal{L}^\perp, i}, \end{aligned}$$

where $\tau_i^2(m)$ denotes the i -th diagonal element of Σ_m . Here we have used the fact that variances are constant within groups, as well as that $\pi_{\mathcal{L}^\perp, i} = 0$ for all $i \in (n_{j-1}^+, n_j^+]$ if $(n_{j-1}^+, n_j^+]$ is disjoint from $I_1(\mathcal{L})$. Also note that the outer sum extends over a nonempty index set since $\text{card}(I_1(\mathcal{L})) \geq 1$ must hold in view of $\dim(\mathcal{L}) < n$. Since the l.h.s. converges to the limit $\bar{\Sigma} \in \mathbb{R}^{n \times n}$, since the r.h.s. is bounded from below in the Loewner order by

$$\|\Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp}\|^{-1} \tau_{n_j^+}^2(m) \sum_{i \in (n_{j-1}^+, n_j^+]} \pi_{\mathcal{L}^\perp, i} \pi'_{\mathcal{L}^\perp, i},$$

for every j appearing in the range of the outer sum, and since $\pi_{\mathcal{L}^\perp, i} \neq 0$ for at least one $i \in (n_{j-1}^+, n_j^+]$ holds when $(n_{j-1}^+, n_j^+] \cap I_1(\mathcal{L}) \neq \emptyset$, it follows that the sequence

$$(\|\Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp}\|^{-1} \tau_{n_j^+}^2(m) : m \in \mathbb{N})$$

is bounded for every j satisfying $(n_{j-1}^+, n_j^+] \cap I_1(\mathcal{L}) \neq \emptyset$. Possibly after passing to a subsequence, we may thus assume that these sequences converge to nonnegative real numbers γ_j for such j . It follows that

$$\begin{aligned} \bar{\Sigma} &= \sum_{j: (n_{j-1}^+, n_j^+] \cap I_1(\mathcal{L}) \neq \emptyset} \gamma_j \sum_{i \in (n_{j-1}^+, n_j^+]} \pi_{\mathcal{L}^\perp, i} \pi'_{\mathcal{L}^\perp, i} \\ &= \sum_{j: (n_{j-1}^+, n_j^+] \cap I_1(\mathcal{L}) \neq \emptyset} \sum_{i \in (n_{j-1}^+, n_j^+]} \gamma_j^{1/2} \pi_{\mathcal{L}^\perp, i} \left(\gamma_j^{1/2} \pi_{\mathcal{L}^\perp, i} \right)'. \end{aligned}$$

Let I be the union of those intervals $(n_{j-1}^+, n_j^+]$ satisfying (i) $(n_{j-1}^+, n_j^+] \cap I_1(\mathcal{L}) \neq \emptyset$ and (ii) $\gamma_j > 0$. Note that I cannot be the empty set as this would imply $\bar{\Sigma} = 0$, which is impossible since it is the limit of a sequence of matrices residing in the unit sphere of $\mathbb{R}^{n \times n}$. Furthermore, by construction, $I \in \mathbf{I}_{(n_1, \dots, n_m)}$ and $I \cap I_1(\mathcal{L}) \neq \emptyset$ hold. Using the fact that $\text{span}(\sum_{l=1}^L A_l A_l') = \text{span}(A_1, \dots, A_L)$ holds for arbitrary real matrices of the same row-dimension, we obtain $\mathcal{S} = \text{span}(\bar{\Sigma}) = \text{span}(\{\pi_{\mathcal{L}^\perp, i} : i \in I\})$ for the before constructed set I . [Note that $\pi_{\mathcal{L}^\perp, i} = 0$ if $i \in (n_{j-1}^+, n_j^+]$ but $i \notin I_1(\mathcal{L})$.] Since \mathcal{S} , being an element of $\mathbb{J}(\mathcal{L}, \mathfrak{C}_{(n_1, \dots, n_m)})$, satisfies $\dim(\mathcal{S}) < n - \dim(\mathcal{L})$, we have established that \mathcal{S} is also an element of (29).

To prove the converse, suppose that \mathcal{S} is an element of (29), i.e., that $\mathcal{S} = \text{span}(\{\pi_{\mathcal{L}^\perp, i} : i \in I\})$ for some $I \in \mathbf{I}_{(n_1, \dots, n_m)}$ with $\emptyset \neq I \cap I_1(\mathcal{L})$ and that $\dim(\mathcal{S}) < n - \dim(\mathcal{L})$ holds. Note that $\text{card}(I) < n$ holds, since otherwise $\mathcal{S} = \mathcal{L}^\perp$ would follow, contradicting $\dim(\mathcal{S}) < n - \dim(\mathcal{L})$. Also note that $\text{card}(I) \geq 1$ as $\emptyset \neq I \cap I_1(\mathcal{L})$. Define diagonal $n \times n$ matrices Σ_m via their diagonal elements

$$\tau_i^2(m) = \begin{cases} (\text{card}(I))^{-1} - \delta_m & \text{if } i \in I \\ (\text{card}(I)/(n - \text{card}(I))) \delta_m & \text{if } i \notin I \end{cases}$$

where $0 < \delta_m < 1/\text{card}(I)$ with $\delta_m \rightarrow 0$ for $m \rightarrow \infty$. Then $\tau_i^2(m) > 0$ as well as $\sum_{i=1}^n \tau_i^2(m) = 1$ hold, and $\tau_{n_{j-1}^+ + 1}^2(m) = \dots = \tau_{n_j^+}^2(m)$ holds for $j = 1, \dots, m$ since $I \in \mathbf{I}_{(n_1, \dots, n_m)}$. That is, Σ_m belongs to $\mathfrak{C}_{(n_1, \dots, n_m)}$. Obviously, Σ_m converges to a diagonal matrix Σ^* with diagonal elements given by

$$\tau_i^{*2} = \begin{cases} (\text{card}(I))^{-1} & \text{if } i \in I \\ 0 & \text{if } i \notin I \end{cases}.$$

Consequently, $\Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp} / \|\Pi_{\mathcal{L}^\perp} \Sigma_m \Pi_{\mathcal{L}^\perp}\|$ converges to $\bar{\Sigma} := \Pi_{\mathcal{L}^\perp} \Sigma^* \Pi_{\mathcal{L}^\perp} / \|\Pi_{\mathcal{L}^\perp} \Sigma^* \Pi_{\mathcal{L}^\perp}\|$, since

$\Pi_{\mathcal{L}^\perp} \Sigma^* \Pi_{\mathcal{L}^\perp} \neq 0$ in view of

$$\Pi_{\mathcal{L}^\perp} \Sigma^* \Pi_{\mathcal{L}^\perp} = \sum_{i=1}^n \tau_i^{*2} \pi_{\mathcal{L}^\perp, i} \pi'_{\mathcal{L}^\perp, i} = (\text{card}(I))^{-1} \sum_{i \in I} \pi_{\mathcal{L}^\perp, i} \pi'_{\mathcal{L}^\perp, i}$$

and the fact that $\emptyset \neq I \cap I_1(\mathcal{L})$ holds and thus $\pi_{\mathcal{L}^\perp, i} \neq 0$ must hold at least for one $i \in I$. Again using $\text{span}(\sum_{l=1}^L A_l A_l') = \text{span}(A_1, \dots, A_L)$ we arrive at

$$\text{span}(\bar{\Sigma}) = \text{span}(\Pi_{\mathcal{L}^\perp} \Sigma^* \Pi_{\mathcal{L}^\perp}) = \text{span} \left((\text{card}(I))^{-1} \sum_{i \in I} \pi_{\mathcal{L}^\perp, i} \pi'_{\mathcal{L}^\perp, i} \right) = \text{span}(\{\pi_{\mathcal{L}^\perp, i} : i \in I\}) = \mathcal{S}.$$

Because we have assumed that $\dim(\mathcal{S}) < n - \dim(\mathcal{L})$ holds, the preceding display shows that $\mathcal{S} \in \mathbb{J}(\mathcal{L}, \mathfrak{C}_{(n_1, \dots, n_m)})$. ■

Remark B.3. Note that $\mathbb{J}(\mathcal{L}, \mathfrak{C}_{(n_1, \dots, n_m)})$ is empty if $m = 1$ (as can be seen directly from the definition of $\mathbb{J}(\mathcal{L}, \mathfrak{C}_{(n_1, \dots, n_m)})$ or from (29)).

Remark B.4. It is easy to see that the concentration spaces of \mathfrak{C}_{Het} in the sense of Preinerstorfer and Pötscher (2016) are precisely given by all spaces of the form $\text{span}(\{e_i(n) : i \in I\})$ where I varies through all subsets of $\{1, \dots, n\}$ that satisfy $0 < \text{card}(I) < n$. More generally, the concentration spaces of $\mathfrak{C}_{(n_1, \dots, n_m)}$ are precisely given by all spaces of the form $\text{span}(\{e_i(n) : i \in I\})$ where $I \in \mathbf{I}_{(n_1, \dots, n_m)}$ satisfies $0 < \text{card}(I) < n$. [In view of Remark 5.1(i) in Pötscher and Preinerstorfer (2018) these results correspond to the case $\dim(\mathcal{L}) = 0$ in the preceding propositions.]

C Appendix: Proofs for Section 4 and Appendix A.1

The facts collected in the subsequent remark will be used in the proofs further below.

Remark C.1. (i) Suppose Assumption 1 holds. Then the test statistic T_{Het} is a non-sphericity corrected F-type test statistic in the sense of Section 5.4 in Preinerstorfer and Pötscher (2016). More precisely, T_{Het} is of the form (28) in Preinerstorfer and Pötscher (2016) and Assumption 5 in the same reference is satisfied with $\check{\beta} = \hat{\beta}$, $\check{\Omega} = \hat{\Omega}_{Het}$, and $N = \emptyset$. Furthermore, the set N^* defined in (27) of Preinerstorfer and Pötscher (2016) satisfies $N^* = \mathbf{B}$. And also Assumptions 6 and 7 of Preinerstorfer and Pötscher (2016) are satisfied. All these claims follow easily in view of Lemma 4.1 in Preinerstorfer and Pötscher (2016), see also the proof of Theorem 4.2 in that reference.

(ii) The test statistic T_{uc} is also a non-sphericity corrected F-type test statistic in the sense of Section 5.4 in Preinerstorfer and Pötscher (2016) (terminology being somewhat unfortunate here as no correction for the non-sphericity is being attempted). More precisely, T_{uc} is of the form (28) in Preinerstorfer and Pötscher (2016) and Assumption 5 in the same reference is satisfied with $\check{\beta} = \hat{\beta}$, $\check{\Omega} = \hat{\sigma}^2 R (X'X)^{-1} R'$, and $N = \emptyset$. Furthermore, the set N^* defined in (27) of Preinerstorfer and Pötscher (2016) satisfies $N^* = \text{span}(X)$. And also Assumptions 6 and 7 of

Preinerstorfer and Pötscher (2016) are satisfied. All these claims are evident (and obviously do not rely on Assumption 1).

(iii) We note that any non-sphericity corrected F-type test statistic (for testing (3)) in the sense of Section 5.4 in Preinerstorfer and Pötscher (2016), i.e., any test statistic T of the form (28) in Preinerstorfer and Pötscher (2016) that also satisfies Assumption 5 in that reference, is invariant under the group $G(\mathfrak{M}_0)$. Furthermore, the associated set N^* defined in (27) of Preinerstorfer and Pötscher (2016) is even invariant under the larger group $G(\mathfrak{M})$. See Sections 5.1 and 5.4 of Preinerstorfer and Pötscher (2016) as well as Lemma 5.16 in Pötscher and Preinerstorfer (2018) for more information.

Proof of Theorem 4.1: We first prove Part (b). We apply Part (b) of Theorem A.1 with $n_j = 1$ for $j = 1, \dots, n = m$ observing that then $\mathfrak{C}_{(n_1, \dots, n_m)} = \mathfrak{C}_{Het}$ and that condition (24) reduces to (13) (exploiting that \mathbf{B} is a finite union of proper linear subspaces as discussed in Lemma 3.1). This establishes (12). The final claim in Part (b) of the theorem follows from Part (b) of Theorem A.1, if we can show that α^* and C^* given there can be written as claimed in Theorem 4.1: To this end we proceed as follows:⁵³ Choose an element μ_0 of \mathfrak{M}_0 . Observe that $I_1(\mathfrak{M}_0^{lin}) \neq \emptyset$ (since $\dim(\mathfrak{M}_0^{lin}) = k - q < n$), and that for every $i \in I_1(\mathfrak{M}_0^{lin})$ the linear space $\mathcal{S}_i = \text{span}(\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n))$ is 1-dimensional (since $\mathcal{S}_i = \{0\}$ is impossible in view of $i \in I_1(\mathfrak{M}_0^{lin})$), and belongs to $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{Het})$ (since $n - k + q > 1 = \dim(\mathcal{S}_i)$ holds) in view of Proposition B.1 in Section B. Since T_{Het} is $G(\mathfrak{M}_0)$ -invariant (Remark C.1 above), it follows that T_{Het} is constant on $(\mu_0 + \mathcal{S}_i) \setminus \{\mu_0\}$, cf. the beginning of the proof of Lemma 5.11 in Pötscher and Preinerstorfer (2018). Hence, \mathcal{S}_i belongs to \mathbb{H} (defined in Lemma 5.11 in Pötscher and Preinerstorfer (2018)) and consequently for C^* as defined in that lemma

$$C^* \geq \max \left\{ T_{Het}(\mu_0 + \Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n)) : i \in I_1(\mathfrak{M}_0^{lin}) \right\}$$

must hold (recall that $\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n) \neq 0$). To prove the opposite inequality, let \mathcal{S} be an arbitrary element of \mathbb{H} , i.e., $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{Het})$ and T_{Het} is $\lambda_{\mu_0 + \mathcal{S}}$ -almost everywhere equal to a constant $C(\mathcal{S})$, say. Then Proposition B.1 in Section B shows that $\mathcal{S}_i \subseteq \mathcal{S}$ holds for some $i \in I_1(\mathfrak{M}_0^{lin})$. Because of Condition (13) we have $\mathcal{S}_i \not\subseteq \mathbf{B}$ since $\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n)$ and $e_i(n)$ differ only by an element of $\mathfrak{M}_0^{lin} \subseteq \text{span}(X)$ and since $\mathbf{B} + \text{span}(X) = \mathbf{B}$. Thus $\mu_0 + \mathcal{S}_i \not\subseteq \mathbf{B}$ by the same argument as $\mu_0 \in \mathfrak{M}_0 \subseteq \text{span}(X)$. We thus can find $s \in \mathcal{S}_i$ such that $\mu_0 + s \notin \mathbf{B}$. Note that $s \neq 0$ must hold, since $\mu_0 \in \mathfrak{M}_0 \subseteq \text{span}(X) \subseteq \mathbf{B}$. In particular, T_{Het} is continuous at $\mu_0 + s$, since $\mu_0 + s \notin \mathbf{B}$. Now, for every open ball A_ε in \mathbb{R}^n with center s and radius $\varepsilon > 0$ we can find an element $a_\varepsilon \in A_\varepsilon \cap \mathcal{S}$ such that $T_{Het}(\mu_0 + a_\varepsilon) = C(\mathcal{S})$. Since $a_\varepsilon \rightarrow s$ for $\varepsilon \rightarrow 0$, it follows that $C(\mathcal{S}) = T_{Het}(\mu_0 + s)$. Since $s \neq 0$ and since T_{Het} is constant on $(\mu_0 + \mathcal{S}_i) \setminus \{\mu_0\}$ as shown before, we can conclude that $C(\mathcal{S}) = T_{Het}(\mu_0 + s) = T_{Het}(\mu_0 + \Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n))$, where we recall that $\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n) \neq 0$. But

⁵³Alternatively, one could base a proof on Lemma C.1 in Pötscher and Preinerstorfer (2019).

this now implies

$$C^* = \max \left\{ T_{Het}(\mu_0 + \Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n)) : i \in I_1(\mathfrak{M}_0^{lin}) \right\}.$$

Using $G(\mathfrak{M}_0)$ -invariance of T_{Het} we conclude that

$$C^* = \max \left\{ T_{Het}(\mu_0 + e_i(n)) : i \in I_1(\mathfrak{M}_0^{lin}) \right\}.$$

The expression for α^* given in the theorem now follows immediately from the expression for α^* given in Part (b) of Theorem A.1.

We next prove Part (a): Apply Part (a) of Theorem A.1 with $n_j = 1$ for $j = 1, \dots, n = m$ observing that then $\mathfrak{C}_{(n_1, \dots, n_m)} = \mathfrak{C}_{Het}$ and that condition (22) reduces to (11) (exploiting that $\text{span}(X)$ is a linear space). This establishes (10). The final claim in Part (a) of the theorem follows similarly as the corresponding claim of Part (b) upon replacing the set B by $\text{span}(X)$ in the argument and by noting that T_{uc} is $G(\mathfrak{M}_0)$ -invariant.

Part (c) follows from Part (c) of Theorem A.1 upon setting $n_j = 1$ for $j = 1, \dots, n = m$ (and upon noting that then the conditions in Theorem A.1 reduce to the conditions of the present theorem). ■

Proof of Proposition 4.5: Follows from Part A.1 of Proposition 5.12 of Pötscher and Preinerstorfer (2018) and the sentence following this proposition. Note that the assumptions of this proposition have been verified in the proof of Theorem 4.1 (see also the proof of Theorem A.1, on which the proof of Theorem 4.1 is based), where it is also shown that the quantity C^* used in Proposition 5.12 of Pötscher and Preinerstorfer (2018) coincides with C^* defined in Theorem 4.1. ■

We note that the result for T_{Het} in Proposition 4.5 can also be obtained from Theorem 4.2 in Preinerstorfer and Pötscher (2016).

Proof of Proposition 4.7: (a) This can be seen as follows (cf. also the discussion on p.302 of Preinerstorfer and Pötscher (2016)): By Remark C.1 above, T_{uc} satisfies the assumptions in Corollary 5.17 in Preinerstorfer and Pötscher (2016) (with $\check{\beta} = \hat{\beta}$, $\check{\Omega}(y) = \hat{\sigma}^2(y)R(X'X)^{-1}R'$, $N = \emptyset$, and $N^* = \text{span}(X)$). Let $e_i(n)$ be one of the standard basis vectors with $i \in I_1(\mathfrak{M}_0^{lin})$ that does belong to $\text{span}(X)$. Set $\mathcal{Z} = \text{span}(e_i(n))$ and note that this is a concentration space of \mathfrak{C}_{Het} , cf. Remark B.4 in Appendix B. The nonnegative definiteness assumption on $\check{\Omega}$ in Part 3 of Corollary 5.17 in Preinerstorfer and Pötscher (2016) is clearly satisfied. We also have $\check{\Omega}(\lambda e_i(n)) = 0$ (since $e_i(n) \in \text{span}(X)$) for every $\lambda \in \mathbb{R}$ and $R\hat{\beta}(\lambda e_i(n)) \neq 0$ for every $\lambda \in \mathbb{R} \setminus \{0\}$ (since $e_i(n) \in \text{span}(X)$ but $e_i(n) \notin \mathfrak{M}_0^{lin}$ in view of $i \in I_1(\mathfrak{M}_0^{lin})$). Part 3 of Corollary 5.17 in Preinerstorfer and Pötscher (2016) then proves the claim for $C > 0$. A fortiori it then also holds for all real C .

(b) This follows for $C > 0$ from Part 3 of Theorem 4.2 in Preinerstorfer and Pötscher (2016) upon observing that a vector $e_i(n)$ satisfying $e_i(n) \in \text{span}(X)$ for some $i \in I_1(\mathfrak{M}_0^{lin})$ clearly satisfies $B(e_i(n)) = 0$ (as $e_i(n) \in \text{span}(X)$) and $R\hat{\beta}(e_i(n)) \neq 0$ (since $e_i(n) \in \text{span}(X)$ but

$e_i(n) \notin \mathfrak{M}_0^{lin}$ in view of $i \in I_1(\mathfrak{M}_0^{lin})$). A fortiori it then also holds for all real C . ■

Proof of Theorem A.1: We first prove Part (b). We wish to apply Part A of Proposition 5.12 of Pötscher and Preinerstorfer (2018) with $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$, $T = T_{Het}$, $\mathcal{L} = \mathfrak{M}_0^{lin}$, and $\mathcal{V} = \{0\}$. First, note that $\dim(\mathfrak{M}_0^{lin}) = k - q < n$. Second, under Assumption 1, T_{Het} is a non-sphericity corrected F-type test with $N^* = \mathbf{B}$, which is a closed $\lambda_{\mathbb{R}^n}$ -null set (see Remarks 3.2 and C.1 as well as Lemma 3.1). Hence, the general assumptions on $T = T_{Het}$, on $N^\dagger = N^* = \mathbf{B}$, on $\mathcal{L} = \mathfrak{M}_0^{lin}$, as well as on \mathcal{V} in Proposition 5.12 of Pötscher and Preinerstorfer (2018) are satisfied in view of Part 1 of Lemma 5.16 in the same reference. [Alternatively, this can be gleaned from Lemma 3.1 and the attending discussion.] Next, observe that condition (24) is equivalent to

$$\text{span} \left(\left\{ \Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n) : i \in (n_{j-1}^+, n_j^+] \right\} \right) \not\subseteq \mathbf{B}$$

for every $j = 1, \dots, m$, such that $(n_{j-1}^+, n_j^+] \cap I_1(\mathfrak{M}_0^{lin}) \neq \emptyset$, since $\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n)$ and $e_i(n)$ differ only by an element of $\mathfrak{M}_0^{lin} \subseteq \text{span}(X)$ and since $\mathbf{B} + \text{span}(X) = \mathbf{B}$ (as noted in Lemma 3.1). In view of Proposition B.2 in Appendix B, this implies that any $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$ is not contained in \mathbf{B} , and thus not in N^\dagger . Using $\mathfrak{M}_0 \subseteq \text{span}(X)$ and $\mathbf{B} + \text{span}(X) = \mathbf{B}$, it follows that $\mu_0 + \mathcal{S} \not\subseteq \mathbf{B} = N^\dagger$ for every $\mu_0 \in \mathfrak{M}_0$. Since $\mu_0 + \mathcal{S}$ is an affine space and $N^\dagger = \mathbf{B}$ is a finite union of proper affine (even linear) spaces under Assumption 1 as discussed in Lemma 3.1, we may conclude (cf. Corollary 5.6 in Pötscher and Preinerstorfer (2018) and its proof) that $\lambda_{\mu_0 + \mathcal{S}}(N^\dagger) = 0$ for every $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$ and every $\mu_0 \in \mathfrak{M}_0$. This completes the verification of the assumptions of Proposition 5.12 in Pötscher and Preinerstorfer (2018) that are not specific to Part A (or Part B) of this proposition. We next verify the assumptions specific to Part A of this proposition: Assumption (a) is satisfied (even for every $C \in \mathbb{R}$) as a consequence of Part 2 of Lemma 5.16 in Pötscher and Preinerstorfer (2018) and of Remark C.1(i) above. And Assumption (b) in Part A follows from Lemma 5.19 of Pötscher and Preinerstorfer (2018), since T_{Het} results as a special case of the test statistics T_{GQ} defined in Section 3.4 of Pötscher and Preinerstorfer (2018) upon choosing $\mathcal{W}_n^* = n^{-1} \text{diag}(d_i)$. Part A of Proposition 5.12 of Pötscher and Preinerstorfer (2018) now immediately delivers claim (23), since $C^* < \infty$ as noted in that proposition. That C^* and α^* do not depend on the choice of $\mu_0 \in \mathfrak{M}_0$ is an immediate consequence of $G(\mathfrak{M}_0)$ -invariance of T_{Het} . Also note that α^* as defined in the theorem coincides with α^* as defined in Proposition 5.12 of Pötscher and Preinerstorfer (2018) in view of $G(\mathfrak{M}_0)$ -invariance of T_{Het} . Positivity of α^* then follows from Part 5 of Lemma 5.15 in Preinerstorfer and Pötscher (2016) in view of Remark C.1(i), noting that $\lambda_{\mathbb{R}^n}$ and $P_{\mu_0, \Sigma}$ are equivalent measures (since $\Sigma \in \mathfrak{C}_{Het}$ is positive definite); cf. Remark 5.13(vi) in Pötscher and Preinerstorfer (2018). In case $\alpha < \alpha^*$, the remaining claim in Part (b) of the theorem, namely that equality can be achieved in (23), follows from the definition of C^* in Lemma 5.11 of Pötscher and Preinerstorfer (2018) and from Part A.2 of Proposition 5.12 of Pötscher and Preinerstorfer (2018) (and the observation immediately following that proposition allowing one to drop the suprema w.r.t. μ_0 and σ^2 , and to set $\sigma^2 = 1$); in case $\alpha = \alpha^* < 1$, it follows from Remarks 5.13(i),(ii) in Pötscher

and Preinerstorfer (2018) using Lemma 5.16 in the same reference.

The proof of Part (a) proceeds along the same lines with some minor differences: Observe that T_{uc} is a non-sphericity corrected F-type test with $N^\dagger = N^* = \text{span}(X)$, which obviously is a closed $\lambda_{\mathbb{R}^n}$ -null set (see Remark C.1(ii)), showing similarly that the general assumptions on $T = T_{uc}$, on $N^\dagger = N^* = \text{span}(X)$, as well as on $\mathcal{L} = \mathfrak{M}_0^{lin}$ in Proposition 5.12 of Pötscher and Preinerstorfer (2018) are again satisfied (with $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$). A similar, even simpler argument as in the proof of Part (b), again shows that condition (22) implies $\lambda_{\mu_0 + \mathcal{S}}(N^\dagger) = 0$ for every $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$ and every $\mu_0 \in \mathfrak{M}_0$, thus completing the verification of the assumptions of Proposition 5.12 of Pötscher and Preinerstorfer (2018) that are not specific to Part A (or Part B) of this proposition. Verification of Assumption (a) in Part A of Proposition 5.12 of Pötscher and Preinerstorfer (2018) proceeds exactly as before. For Assumption (b) we now use Lemma 5.19(iii) of Pötscher and Preinerstorfer (2018), since T_{uc} results as a special case of the test statistics $T_{E,W}$ defined in Section 3 of Pötscher and Preinerstorfer (2018) upon choosing W as $n(n-k)^{-1}I_n$. Part A of Proposition 5.12 of Pötscher and Preinerstorfer (2018) then delivers the claim (21), again since $C^* < \infty$ as noted in that proposition. Again, $G(\mathfrak{M}_0)$ -invariance of T_{uc} implies that C^* and α^* do not depend on the choice of $\mu_0 \in \mathfrak{M}_0$, and that α^* as defined in the theorem coincides with α^* as defined in Proposition 5.12 of Pötscher and Preinerstorfer (2018). Positivity of α^* follows exactly as before making now use of Remark C.1(ii). The remaining claim in Part (a) is proved completely analogous as the corresponding claim in Part (b).

We finally prove Part (c): The first claim follows from Remark 5.10 and Lemma 5.16 in Pötscher and Preinerstorfer (2018) combined with Remark C.1 above. The second claim is obvious. ■

Proof of Proposition A.2: (a) This follows from Part 3 of Corollary 5.17 in Preinerstorfer and Pötscher (2016): As shown in the proof of Proposition 4.7(a) T_{uc} satisfies the assumptions of this corollary (with $\check{\beta} = \hat{\beta}$, $\check{\Omega}(y) = \hat{\sigma}^2(y)R(X'X)^{-1}R'$, $N = \emptyset$, and $N^* = \text{span}(X)$). Set now $\mathcal{Z} = \text{span}(\{e_i(n) : i \in (n_{j-1}^+, n_j^+]\})$, where j is such that $(n_{j-1}^+, n_j^+] \cap I_1(\mathfrak{M}_0^{lin}) \neq \emptyset$ and $\mathcal{Z} \subseteq \text{span}(X)$ hold. Note that \mathcal{Z} is not contained in \mathfrak{M}_0^{lin} by construction. Observe that \mathcal{Z} is a concentration space of $\mathfrak{C}_{(n_1, \dots, n_m)}$ in view of Remark B.4 in Appendix B (note that $\text{card}((n_{j-1}^+, n_j^+]) < n$ must hold in view of $\mathcal{Z} \subseteq \text{span}(X)$ and $k < n$, while $0 < \text{card}((n_{j-1}^+, n_j^+])$ is obvious). The nonnegative definiteness assumption on $\check{\Omega}$ in Part 3 of Corollary 5.17 in Preinerstorfer and Pötscher (2016) is clearly satisfied. Obviously $\check{\Omega}(z) = 0$ holds for every $z \in \mathcal{Z}$ since $\mathcal{Z} \subseteq \text{span}(X)$. It remains to establish that $R\hat{\beta}(z) \neq 0$ holds $\lambda_{\mathcal{Z}}$ -everywhere: Clearly, $R\hat{\beta}(z) = 0$ for $z \in \mathcal{Z}$ occurs precisely for $z \in \mathcal{Z} \cap \mathfrak{M}_0^{lin}$ since $\mathcal{Z} \subseteq \text{span}(X)$. But $\mathcal{Z} \cap \mathfrak{M}_0^{lin}$ is a $\lambda_{\mathcal{Z}}$ -null set in view of the fact that \mathcal{Z} is not contained in \mathfrak{M}_0^{lin} as noted before (and hence $\mathcal{Z} \cap \mathfrak{M}_0^{lin}$ is a proper linear subspace of \mathcal{Z}). Part 3 of Corollary 5.17 in Preinerstorfer and Pötscher (2016) then proves the claim for $C > 0$. A fortiori it then also holds for all real C .

(b) This follows in the same way as Part (a) by applying Part 3 of Corollary 5.17 in Preinerstorfer and Pötscher (2016) now to T_{Het} (with $\check{\beta} = \hat{\beta}$, $\check{\Omega} = \hat{\Omega}_{Het}$, $N = \emptyset$, and $N^* = B$). ■

We note that Propositions 4.7 and A.2 could also be proved by making use of Theorem 3.1 in Pötscher and Preinerstorfer (2019).

Remark C.2. (i) Condition (11) ((13), respectively) in Theorem 4.1 can equivalently be written as $\text{span}(\{\pi_{(\mathfrak{M}_0^{lin})^\perp, i}\}) \not\subseteq \text{span}(X)$ ($\not\subseteq \mathbf{B}$, respectively) for every $i \in I_1(\mathfrak{M}_0^{lin})$ as discussed in the proof. Since the spaces $\text{span}(\{\pi_{(\mathfrak{M}_0^{lin})^\perp, i}\})$ are one-dimensional for $i \in I_1(\mathfrak{M}_0^{lin})$ and since $1 < n - k + q = n - \dim(\mathfrak{M}_0^{lin})$, it follows that these spaces are necessarily elements of $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{Het})$; in fact, they are precisely the minimal elements of $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{Het})$ w.r.t. the order induced by inclusion.

(ii) Condition (22) ((24), respectively) in Theorem A.1 can equivalently be written as

$$\text{span}(\{\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n) : i \in (n_{j-1}^+, n_j^+]\}) \not\subseteq \text{span}(X) \quad (\not\subseteq \mathbf{B}, \text{ respectively})$$

for every $j = 1, \dots, m$ with $(n_{j-1}^+, n_j^+] \cap I_1(\mathfrak{M}_0^{lin}) \neq \emptyset$ as discussed in the proof. However, in this more general case, it can happen that such a space appearing on the l.h.s. of the non-inclusion relation has a dimension not smaller than $n - \dim(\mathfrak{M}_0^{lin})$, and hence is not a member of $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$. In light of the general results in Pötscher and Preinerstorfer (2018) (e.g., Corollary 5.6) one may wonder if requiring the non-inclusion condition in (22) (24, respectively) for such spaces does not add an unnecessary restriction. However, this is not so as this non-inclusion is easily seen to be automatically satisfied for such spaces.⁵⁴ Furthermore, the collection of all spaces of the form $\text{span}(\{\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n) : i \in (n_{j-1}^+, n_j^+]\})$ for $j = 1, \dots, m$, such that $(n_{j-1}^+, n_j^+] \cap I_1(\mathfrak{M}_0^{lin}) \neq \emptyset$ and such that the dimension of these spaces is smaller than $n - \dim(\mathfrak{M}_0^{lin})$ is precisely the collection of minimal elements of $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$ w.r.t. the order induced by inclusion. [Note that $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$ may be empty.]

Proposition C.3. *Suppose we are in the setting of Example 4.5 with $n_j \geq 2$ for all j . Then T_{Het} is size controllable over \mathfrak{C}_{Het} , i.e., (12) holds for every $0 < \alpha < 1$.*

Proof: Note that \mathbf{B} is a subset of

$$\mathbf{S} := \{y \in \mathbb{R}^n : \hat{u}_i(y) = 0 \text{ for some } i = 1, \dots, n\},$$

and that \mathbf{S} is a $\lambda_{\mathbb{R}^n}$ -null set, as it is a finite union of $\lambda_{\mathbb{R}^n}$ -null sets (since $e_i(n) \notin \text{span}(X)$ in view of $n_j \geq 2$ for all j). Also note that $S_j > 0$ holds for $y \notin \mathbf{S}$. Now, for $y \notin \mathbf{S}$, by the Sherman-Morrison formula, the inverse of $S_1 \mu' + \text{diag}(S_2, \dots, S_k)$ equals

$$\text{diag}(S_2^{-1}, \dots, S_k^{-1}) - \text{diag}(S_2^{-1}, \dots, S_k^{-1}) \mu' \text{diag}(S_2^{-1}, \dots, S_k^{-1}) / \sum_{j=1}^k 1/S_j.$$

⁵⁴Note that any such space is necessarily equal to $(\mathfrak{M}_0^{lin})^\perp$. If now $(\mathfrak{M}_0^{lin})^\perp$ were contained in $\text{span}(X)$ (\mathbf{B} , respectively), then \mathbb{R}^n would also have to be contained in $\text{span}(X)$ (\mathbf{B} , respectively), since \mathbb{R}^n can be written as the direct sum of $(\mathfrak{M}_0^{lin})^\perp$ and \mathfrak{M}_0^{lin} and since $\text{span}(X)$ (\mathbf{B} , respectively) are invariant under addition of elements of \mathfrak{M}_0^{lin} . However, $\text{span}(X)$ is a proper subspace of \mathbb{R}^n (since we always assume $k < n$) and \mathbf{B} is a finite union of proper linear subspaces of \mathbb{R}^n under Assumption 1. This gives a contradiction.

We may thus write

$$T_{Het}(y) = \sum_{j=2}^k \frac{(\bar{y}_{(1)} - \bar{y}_{(j)})^2}{S_j} - \left[\sum_{j=2}^k \frac{\bar{y}_{(1)} - \bar{y}_{(j)}}{S_j} \right]^2 / \sum_{j=1}^k 1/S_j \quad \text{for every } y \notin \mathbf{S}. \quad (30)$$

As noted in Remark 3.6, for any invertible $q \times q$ -dimensional matrix A , the test statistic T_{Het} based on R and the analogous test statistic, but computed with AR instead of R , coincide everywhere (note $r = 0$). We apply this observation in the following way: fix $l \in \{2, \dots, k\}$, and choose A with l -th column $(-1, \dots, -1)'$, l -th row $(0, \dots, 0, -1, 0, \dots, 0)$, and such that after deleting the l -th column and the l -th row we obtain I_{q-1} . Then

$$AR = RP_l,$$

where P_l is the $k \times k$ permutation matrix that interchanges the first and l -th coordinate (and keeps all other coordinates fixed). By a similar computation as the one that led to the expression in (30), but now with RP_l in place of R , we can now conclude that for every $l \in \{1, \dots, k\}$ we have

$$T_{Het}(y) = \sum_{j=1, j \neq l}^k \frac{(\bar{y}_{(l)} - \bar{y}_{(j)})^2}{S_j} - \left[\sum_{j=1, j \neq l}^k \frac{\bar{y}_{(l)} - \bar{y}_{(j)}}{S_j} \right]^2 / \sum_{j=1}^k 1/S_j \quad \text{for every } y \notin \mathbf{S}.$$

For $y \notin \mathbf{S}$ we may thus upper bound $T_{Het}(y)$ by $\sum_{j \neq l}^k (\bar{y}_l - \bar{y}_j)^2 / S_j$, and we are free to choose l . Setting $l = l(y) \in \arg \min_{j=1, \dots, k} S_j$, the upper bound for $T_{Het}(y)$ just derived, together with $S_j \geq (S_j + S_l)/2 > 0$, gives for $y \notin \mathbf{S}$

$$\begin{aligned} T_{Het}(y) &\leq \sum_{j=1, j \neq l}^k \frac{(\bar{y}_{(l)} - \bar{y}_{(j)})^2}{S_j} \leq 2 \sum_{j=1, j \neq l}^k \frac{(\bar{y}_{(l)} - \bar{y}_{(j)})^2}{S_j + S_l} \\ &\leq 2 \sum_{i, j=1, i \neq j}^k \frac{(\bar{y}_{(i)} - \bar{y}_{(j)})^2}{S_i + S_j} = 2 \sum_{i, j=1, i \neq j}^k T_{i,j}(y), \end{aligned}$$

where $T_{i,j}(y) = (\bar{y}_{(i)} - \bar{y}_{(j)})^2 / (S_i + S_j)$. Note that the quantity to the far right does not depend on our particular choice of l . For $y \in \mathbf{S}$, define $T_{i,j}$ by the same formula as long as $S_i + S_j > 0$ and as $T_{i,j} = 0$ else. Since \mathbf{S} is a $\lambda_{\mathbb{R}^n}$ -null set, we have for any $C > 0$

$$\sup_{\mu_0 \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(T_{Het} \geq C) \leq \sum_{i, j=1, i \neq j}^k \sup_{\mu_0 \in \mathfrak{M}_0} \sup_{0 < \sigma^2 < \infty} \sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \sigma^2 \Sigma}(T_{i,j} \geq C/2(k^2 - k)).$$

Now observe that $T_{i,j}$ depends only on the coordinates of y corresponding to groups i and j and furthermore coincides with the test statistic of the form (4) for a two sample mean comparison as considered in Example 4.4 (with sample size being equal to $n_i + n_j$). A simple argument then

shows that the terms in the sum on the r.h.s of the preceding display can be rewritten as the sizes of the test statistic (4) as considered in Example 4.4 with sample size now being given by $n_i + n_j$. Hence, all these terms can be made arbitrarily small by choosing C large enough by what has been established in Example 4.4. ■

We provide here a further example, where the sufficient condition of Part (b) of Theorem 4.1 fails, but size control is possible.

Example C.1. Suppose we are given $k \geq 2$ integers n_j describing group sizes satisfying $n_1 \geq 2$ and $n_j \geq 1$ for $j \geq 2$. Sample size is $n = \sum_{j=1}^k n_j$. Clearly $k < n$ is then satisfied. The regressors x_{ti} indicate group membership, i.e., they satisfy $x_{ti} = 1$ for $\sum_{j=1}^{i-1} n_j < t \leq \sum_{j=1}^i n_j$ and $x_{ti} = 0$ otherwise. The heteroskedasticity model is again given by \mathfrak{C}_{Het} . Let $R = (1, 0, \dots, 0)$, i.e., the coefficient of the first regressor is subject to test. Then $I_0(\mathfrak{M}_0^{lin}) = \{\sum_{l=1}^j n_l : n_j = 1, j = 2, \dots, k\}$. With regard to T_{uc} we immediately see that $e_i(n) \notin \text{span}(X)$ for $i \in I_1(\mathfrak{M}_0^{lin})$ holds, and thus the sufficient condition (11) for size control of T_{uc} is satisfied. Turning to T_{Het} , observe that Assumption 1 is satisfied as is easily seen. Furthermore, it is not difficult to see that $\mathbf{B} = \{y \in \mathbb{R}^n : y_1 = \dots = y_{n_1}\}$. Note that $\text{span}(X) \subseteq \mathbf{B}$, but $\mathbf{B} \neq \text{span}(X)$, except if $n_j = 1$ for all $j \geq 2$ holds. In the latter case it is then easy to see that $e_i(n) \notin \text{span}(X) = \mathbf{B}$ for every $i \in I_1(\mathfrak{M}_0^{lin})$ holds, and thus the sufficient condition (13) for size control of T_{Het} is satisfied. But if $n_j > 1$ for some $j \geq 2$ holds, then for any index i satisfying $\sum_{l=1}^{j-1} n_l < i \leq \sum_{l=1}^j n_l$ we have $i \in I_1(\mathfrak{M}_0^{lin})$ as well as $e_i(n) \in \mathbf{B}$. Consequently, the sufficient condition (13) for size control of T_{Het} is *not* satisfied and hence Theorem 4.1 does not inform us about size controllability of T_{Het} in this case. However, the following argument shows that size control for T_{Het} is possible also in this case: The test statistic T_{Het} for the given problem coincides with a corresponding test statistic (again of the form (6) for an appropriate choice of d_i 's) in the “reduced” problem that one obtains by throwing away all data points for $t > n_1$ and by also deleting all regressors from the regression model but the first one. This leads one to the heteroskedastic location model discussed in Example 4.3 albeit with sample size reduced to n_1 . It is now not difficult to see that the size of T_{Het} in the original formulation of the problem coincides with the size of the corresponding test statistic in the “reduced” problem, which – in light of the discussion in Example 4.3 – shows that size control for T_{Het} in the original problem is possible also in the case where $n_j > 1$ for some $j \geq 2$ holds. [If $n_1 = 1$ and if $n_j \geq 2$ for some j , condition (11) in Theorem 4.1(a) is violated, implying – in view of Proposition 4.7 – that the size of the rejection region $\{T_{uc} > C\}$ is 1 for every choice of C ; and that the test statistic T_{Het} is identically zero (since Assumption 1 is violated and, in fact, $\hat{\Omega}_{Het}$ is identically zero). The case where all n_j are equal to 1 even falls outside of our framework since we always require $n > k$.]

Remark C.4. Alternatively to the argument given in the above example for the case where $n_j > 1$ for some $j \geq 2$ holds, size controllability of T_{Het} can also be established by the following reasoning: Keep the sample of size n , but replace the regressors $x_{\cdot i}$ for $2 \leq i \leq k$ by new regressors given by the standard basis vectors $e_j(n)$ for $j > n_1$ (the number of regressors now being $k^* = n - n_1 + 1 < n$ and $R = (1, 0, \dots, 0)$ now being $1 \times k^*$). Then one observes that

(i) this does not affect the test statistic, (ii) makes the set \mathfrak{M}_0 at most larger, and (ii) in the new model the sufficient condition (13) is now satisfied (as in the new model $n_j = 1$ holds for $j > n_1$). Hence, size control (even over the larger \mathfrak{M}_0) follows. A third possibility to establish the size-controllability result is to observe that the test statistic T_{Het} as well as the set \mathbf{B} in the original model are – additional to being $G(\mathfrak{M}_0)$ -invariant – also invariant w.r.t. addition of the elements $e_i(n)$ for $i > n_1$ and then to appeal to a generalization of Theorem 4.1 that exploits this additional invariance and provides sufficient conditions for size control that can be seen to be satisfied in the model considered in this example. Such a generalization of Theorem 4.1, which we refrain from stating, can be obtained from the general size control results presented in Pötscher and Preinerstorfer (2018).

Remark C.5. Example C.1 is an instance of the following observation: Suppose X is block-diagonal of rank k with blocks X_1 and X_2 where X_i is $n_i \times k_i$ with $n_1 + n_2 = n$ and $k_1 + k_2 = k$. Assume $k_1 < n_1$ (which entails $k < n$). Assume that the $q \times k$ restriction matrix R is of rank q and has the form $R = (R_1 : 0)$ with R_1 of dimension $q \times k_1$. The heteroskedasticity model is given by \mathfrak{C}_{Het} . Then, using the same reasoning as in Example C.1, we see that the question of size control of T_{Het} is equivalent to the question of size control of the corresponding test statistic in the “reduced” problem where one considers the regression model with regressor matrix equal to X_1 using only observations with $t \leq n_1$ (and as heteroskedasticity model the analogue of \mathfrak{C}_{Het} for sample size n_1). As Example C.1 has shown, it is possible that the sufficient conditions for size control of T_{Het} are violated in the “original” problem, while at the same time the sufficient conditions may be satisfied in the “reduced” problem. Alternatively, one can argue similarly as in Remark C.4.

D Appendix: Proofs for Section 5 and Appendix A.2

Lemma D.1. (a) Let \mathcal{S} be a linear subspace of \mathbb{R}^n and μ an element of \mathbb{R}^n such that \tilde{T}_{uc} restricted to $\mu + \mathcal{S}$ is not equal to a constant $\lambda_{\mu+\mathcal{S}}$ -almost everywhere. Then $\lambda_{\mu+\mathcal{S}}(\tilde{T}_{uc} = C) = 0$ holds for every $C \in \mathbb{R}$.

(b) $\lambda_{\mathbb{R}^n}(\tilde{T}_{uc} = C) = 0$ holds for every $C \in \mathbb{R}$.

(c) Let \mathcal{S} be a linear subspace of \mathbb{R}^n and μ an element of \mathbb{R}^n such that \tilde{T}_{Het} restricted to $\mu + \mathcal{S}$ is not equal to a constant $\lambda_{\mu+\mathcal{S}}$ -almost everywhere. Then $\lambda_{\mu+\mathcal{S}}(\tilde{T}_{Het} = C) = 0$ holds for every $C \in \mathbb{R}$.

(d) Suppose Assumption 2 holds and \tilde{T}_{Het} is not constant on $\mathbb{R}^n \setminus \tilde{\mathbf{B}}$. Then $\lambda_{\mathbb{R}^n}(\tilde{T}_{Het} = C) = 0$ holds for every $C \in \mathbb{R}$.

Proof: (a) Since \tilde{T}_{uc} is constant on \mathfrak{M}_0 by definition, it follows that $\mu + \mathcal{S} \not\subseteq \mathfrak{M}_0$ must hold, and hence \mathfrak{M}_0 is a $\lambda_{\mu+\mathcal{S}}$ -null set (cf. the argument in Remark 5.9(i) in Pötscher and Preinerstorfer (2018)). Consequently, \tilde{T}_{uc} restricted to $(\mu + \mathcal{S}) \setminus \mathfrak{M}_0$ is not constant. Suppose now there exists a $C \in \mathbb{R}$ so that $\lambda_{\mu+\mathcal{S}}(\{y \in \mathbb{R}^n : \tilde{T}_{uc}(y) = C\}) > 0$. Then, since \mathfrak{M}_0 is a $\lambda_{\mu+\mathcal{S}}$ -null set as just shown, it follows that even $\lambda_{\mu+\mathcal{S}}(\{y \in \mathbb{R}^n \setminus \mathfrak{M}_0 : \tilde{T}_{uc}(y) = C\}) > 0$ must hold,

which can be written as $\lambda_{\mu+\mathcal{S}}(\{y \in \mathbb{R}^n \setminus \mathfrak{M}_0 : p(y) = 0\}) > 0$, with the multivariate polynomial p given by $p(y) = (R\hat{\beta}(y) - r)'(R(X'X)^{-1}R')^{-1}(R\hat{\beta}(y) - r) - C\tilde{\sigma}^2(y)$. This implies that p restricted to $\mu + \mathcal{S}$ vanishes on a set of positive $\lambda_{\mu+\mathcal{S}}$ -measure. Since p restricted to $\mu + \mathcal{S}$ can clearly be expressed as a polynomial in coordinates parameterizing the affine space $\mu + \mathcal{S}$, it follows that p vanishes identically on $\mu + \mathcal{S}$. But this implies that \tilde{T}_{uc} restricted to $(\mu + \mathcal{S}) \setminus \mathfrak{M}_0$ is constant equal to C , a contradiction (as \mathfrak{M}_0 is a $\lambda_{\mu+\mathcal{S}}$ -null set).

(b) Follows from Part (a) upon choosing $\mathcal{S} = \mathbb{R}^n$, if we can show that \tilde{T}_{uc} is not $\lambda_{\mathbb{R}^n}$ -almost everywhere constant. Given that \tilde{T}_{uc} is continuous on $\mathbb{R}^n \setminus \mathfrak{M}_0$ (the complement of a proper affine subspace), it suffices to show that \tilde{T}_{uc} is not constant on $\mathbb{R}^n \setminus \mathfrak{M}_0$. To this end consider first $y = X\beta$ with $R\beta - r \neq 0$ (such a β obviously exists). Observe that $\tilde{\sigma}^2(y) \neq 0$ as $y \notin \mathfrak{M}_0$ and that $R\hat{\beta}(y) - r = R\beta - r \neq 0$. Hence, $\tilde{T}_{uc}(y) \neq 0$ for this choice of y . Next, choose $y = X\beta + w$, where $R\beta - r = 0$ (such a β obviously exists) and where $w \neq 0$ is orthogonal to $\text{span}(X)$ (which is possible since $k < n$ is always maintained). Then $\hat{\beta}(y) = \beta = \tilde{\beta}(y)$, implying $R\hat{\beta}(y) - r = R\beta - r = 0$ and $\tilde{\sigma}^2(y) = w'w/(n - (k - q)) \neq 0$. Note that $y \notin \mathfrak{M}_0$. It follows that $\tilde{T}_{uc}(y) = 0$ holds for this choice of y . This establishes non-constancy of \tilde{T}_{uc} on $\mathbb{R}^n \setminus \mathfrak{M}_0$.

(c) Completely analogous to the proof of Part (a) except that \tilde{T}_{uc} and \mathfrak{M}_0 are replaced by \tilde{T}_{Het} and $\tilde{\mathfrak{B}}$, respectively, and that p now takes the form $p(y) = (R\hat{\beta}(y) - r)' \text{adj}(\tilde{\Omega}_{Het}(y))(R\hat{\beta}(y) - r) - C \det(\tilde{\Omega}_{Het}(y))$, where $\text{adj}(\cdot)$ denotes the adjoint of the square matrix indicated, with the convention that the adjoint of a 1×1 dimensional matrix equals one. [We note that under the assumptions for Part (c) the set $\tilde{\mathfrak{B}}$ cannot coincide with \mathbb{R}^n (since otherwise \tilde{T}_{Het} would be constant equal to zero), and thus Assumption 2 must hold.]

(d) Follows from Part (c) upon choosing $\mathcal{S} = \mathbb{R}^n$, if we can show that \tilde{T}_{Het} is not $\lambda_{\mathbb{R}^n}$ -almost everywhere constant. Given that \tilde{T}_{Het} is continuous on $\mathbb{R}^n \setminus \tilde{\mathfrak{B}}$ (the complement of a finite union of proper affine subspaces by Lemma 3.3), this follows from the assumed non-constancy on $\mathbb{R}^n \setminus \tilde{\mathfrak{B}}$. ■

Remark D.2. The additional assumption that \tilde{T}_{Het} is not constant on $\mathbb{R}^n \setminus \tilde{\mathfrak{B}}$ in Part (d) of the preceding lemma can not be dropped as can be seen from the following example: Consider the case where $k = q = 1$, $R = 1$, $r = 0$, the regressor is given by $e_1(n)$, and the constants \tilde{d}_i satisfy $\tilde{d}_i = 1$ for all i . Then $\mathfrak{M}_0 = \mathfrak{M}_0^{lin} = \{0\}$, Assumption 2 is satisfied, and $\tilde{\mathfrak{B}} = \text{span}(e_1(n))^\perp$. Furthermore, $\tilde{T}_{Het}(y) = 1$ for every $y \in \mathbb{R}^n \setminus \tilde{\mathfrak{B}}$. As a point of interest we note that \tilde{T}_{Het} is trivially size controllable for every $0 < \alpha < 1$, but that the condition (17) for size controllability is violated since $e_j(n) \in \tilde{\mathfrak{B}}$ for $j > 1$. [Of course, neither a smallest size-controlling critical value exists (when considering rejection regions of the form $\{\tilde{T}_{Het} \geq C\}$) nor can exact size controllability be achieved for $0 < \alpha < 1$.]

Lemma D.3. *The rejection probabilities $P_{\mu_0, \sigma^2 \Sigma}(\tilde{T}_{uc} \geq C)$ as well as $P_{\mu_0, \sigma^2 \Sigma}(\tilde{T}_{Het} \geq C)$ for $\mu_0 \in \mathfrak{M}_0$, $\sigma^2 \in (0, \infty)$, $\Sigma \in \mathfrak{C}_{Het}$, do not depend on r . [It is understood here that the constants \tilde{d}_i appearing in the definition of \tilde{T}_{Het} have been chosen independently of the value of r .]*

Proof: Fix $\mu_0 \in \mathfrak{M}_0$. Observe that $\tilde{T}_{Het}(y) = \tilde{T}_{Het}^0(y - \mu_0)$, where

$$\tilde{T}_{Het}^0(z) = \begin{cases} (R\hat{\beta}(z))' \left(\tilde{\Omega}_{Het}^0(z) \right)^{-1} (R\hat{\beta}(z)) & \text{if } \text{rank } \tilde{B}^0(z) = q, \\ 0 & \text{if } \text{rank } \tilde{B}^0(z) < q, \end{cases}$$

where

$$\tilde{\Omega}_{Het}^0(z) = R(X'X)^{-1}X' \text{diag} \left(\tilde{d}_1(\tilde{u}_1^0(z))^2, \dots, \tilde{d}_n(\tilde{u}_n^0(z))^2 \right) X(X'X)^{-1}R',$$

where $\tilde{u}^0(z) = \Pi_{(\mathfrak{M}_0^{lin})^\perp} z$, and where $\tilde{B}^0(z) = R(X'X)^{-1}X' \text{diag}(e'_1(n)\Pi_{(\mathfrak{M}_0^{lin})^\perp}(z), \dots, e'_n(n)\Pi_{(\mathfrak{M}_0^{lin})^\perp}(z))$. Here we have made use of (8) and the fact that $\tilde{u}(y) = \tilde{u}^0(y - \mu_0)$. Now

$$P_{\mu_0, \sigma^2 \Sigma}(\tilde{T}_{Het}(y) \geq C) = P_{\mu_0, \sigma^2 \Sigma}(\tilde{T}_{Het}^0(y - \mu_0) \geq C) = P_{0, \sigma^2 \Sigma}(\tilde{T}_{Het}^0(z) \geq C)$$

and the far right-hand side does not depend on r as \tilde{T}_{Het}^0 does not depend on r . The proof for \tilde{T}_{uc} is completely analogous, noting that $\tilde{T}_{uc}(y) = \tilde{T}_{uc}^0(y - \mu_0)$, where

$$\tilde{T}_{uc}^0(z) = \begin{cases} (R\hat{\beta}(z))' ((\tilde{\sigma}^0(z))^2 R(X'X)^{-1}R')^{-1} (R\hat{\beta}(z)) & \text{if } z \notin \mathfrak{M}_0^{lin}, \\ 0 & \text{if } z \in \mathfrak{M}_0^{lin}, \end{cases}$$

and where $(\tilde{\sigma}^0(z))^2 = (\tilde{u}^0(z))' \tilde{u}^0(z) / (n - (k - q))$. ■

Proof of Theorem 5.1: We first prove Part (b). We apply Part (b) of Theorem A.5 with $n_j = 1$ for $j = 1, \dots, n = m$ observing that then $\mathfrak{C}_{(n_1, \dots, n_m)} = \mathfrak{C}_{Het}$ and that condition (27) reduces to (17) (exploiting that $\tilde{\mathfrak{B}} - \mu_0$ is a finite union of proper linear subspaces as discussed in Remark 3.4). This establishes (16). The final claim in Part (b) of the theorem follows from Part (b) of Theorem A.5, if we can show that C^* given there can be written as claimed in Theorem 5.1: To this end we proceed as follows:⁵⁵ Choose an element μ_0 of \mathfrak{M}_0 . Observe that $I_1(\mathfrak{M}_0^{lin}) \neq \emptyset$ (since $\dim(\mathfrak{M}_0^{lin}) = k - q < n$), and that for every $i \in I_1(\mathfrak{M}_0^{lin})$ the linear space $\mathcal{S}_i = \text{span}(\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n))$ is 1-dimensional (since $\mathcal{S}_i = \{0\}$ is impossible in view of $i \in I_1(\mathfrak{M}_0^{lin})$), and belongs to $\mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{Het})$ (since $n - k + q > 1 = \dim(\mathcal{S}_i)$ holds) in view of Proposition B.1 in Section B. Since \tilde{T}_{Het} is $G(\mathfrak{M}_0)$ -invariant (Remark 3.5), it follows that \tilde{T}_{Het} is constant on $(\mu_0 + \mathcal{S}_i) \setminus \{\mu_0\}$, cf. the beginning of the proof of Lemma 5.11 in Pötscher and Preinerstorfer (2018). Hence, \mathcal{S}_i belongs to \mathbb{H} (defined in Lemma 5.11 in Pötscher and Preinerstorfer (2018)) and consequently for C^* as defined in that lemma

$$C^* \geq \max \left\{ \tilde{T}_{Het}(\mu_0 + \Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n)) : i \in I_1(\mathfrak{M}_0^{lin}) \right\}$$

must hold. To prove the opposite inequality, let \mathcal{S} be an arbitrary element of \mathbb{H} , i.e., $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{Het})$ and \tilde{T}_{Het} is $\lambda_{\mu_0 + \mathcal{S}}$ -almost everywhere equal to a constant $C(\mathcal{S})$, say. Then Proposition B.1 in Section B shows that $\mathcal{S}_i \subseteq \mathcal{S}$ holds for some $i \in I_1(\mathfrak{M}_0^{lin})$. Because of Condition

⁵⁵Alternatively, one could base a proof on Lemma C.1 in Pötscher and Preinerstorfer (2019).

(17) we have $\mu_0 + \mathcal{S}_i \not\subseteq \tilde{\mathcal{B}}$ since $\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n)$ and $e_i(n)$ differ only by an element of \mathfrak{M}_0^{lin} and since $\tilde{\mathcal{B}} + \mathfrak{M}_0^{lin} = \tilde{\mathcal{B}}$. We thus can find $s \in \mathcal{S}_i$ such that $\mu_0 + s \notin \tilde{\mathcal{B}}$. Note that $s \neq 0$ must hold, since $\mu_0 \in \mathfrak{M}_0 \subseteq \tilde{\mathcal{B}}$ (see Lemma 3.3). In particular, \tilde{T}_{Het} is continuous at $\mu_0 + s$, since $\mu_0 + s \notin \tilde{\mathcal{B}}$. Now, for every open ball A_ε in \mathbb{R}^n with center s and radius $\varepsilon > 0$ we can find an element $a_\varepsilon \in A_\varepsilon \cap \mathcal{S}$ such that $\tilde{T}_{Het}(\mu_0 + a_\varepsilon) = C(\mathcal{S})$. Since $a_\varepsilon \rightarrow s$ for $\varepsilon \rightarrow 0$, it follows that $C(\mathcal{S}) = \tilde{T}_{Het}(\mu_0 + s)$. Since $s \neq 0$ and since \tilde{T}_{Het} is constant on $(\mu_0 + \mathcal{S}_i) \setminus \{\mu_0\}$ as shown before, we can conclude that $C(\mathcal{S}) = \tilde{T}_{Het}(\mu_0 + s) = \tilde{T}_{Het}(\mu_0 + \Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n))$, where we recall that $\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n) \neq 0$. But this now implies

$$C^* = \max \left\{ \tilde{T}_{Het}(\mu_0 + \Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n)) : i \in I_1(\mathfrak{M}_0^{lin}) \right\}.$$

Using $G(\mathfrak{M}_0)$ -invariance of \tilde{T}_{Het} we conclude that

$$C^* = \max \left\{ \tilde{T}_{Het}(\mu_0 + e_i(n)) : i \in I_1(\mathfrak{M}_0^{lin}) \right\}.$$

We next prove Part (a): Apply Part (a) of Theorem A.5 with $n_j = 1$ for $j = 1, \dots, n = m$, observing that then $\mathfrak{C}_{(n_1, \dots, n_m)} = \mathfrak{C}_{Het}$. This establishes (15).⁵⁶ The final claim in Part (a) of the theorem follows similarly as the corresponding claim of Part (b) upon replacing the set $\tilde{\mathcal{B}}$ by \mathfrak{M}_0 in the argument, by noting that \tilde{T}_{uc} is $G(\mathfrak{M}_0)$ -invariant, and that $\mu_0 + \mathcal{S}_i \not\subseteq \mathfrak{M}_0$ holds because of $i \in I_1(\mathfrak{M}_0^{lin})$.

Part (c) follows from Part (c) of Theorem A.5 upon setting $n_j = 1$ for $j = 1, \dots, n = m$ (and upon noting that then the conditions in Theorem A.5 reduce to the conditions of the present theorem). ■

Proof of Proposition 5.4: Follows from Part A.1 of Proposition 5.12 of Pötscher and Preinerstorfer (2018) and the sentence following this proposition. Note that the assumptions of this proposition have been verified in the proof of Theorem 5.1 (see also the proof of Theorem A.5, on which the proof of Theorem 5.1 is based), where it is also shown that the quantity C^* used in Proposition 5.12 of Pötscher and Preinerstorfer (2018) coincides with C^* defined in Theorem 5.1. ■

Proof of Theorem 5.8: From the definition of C^* we see that C^* is nonnegative and finite. Let C be arbitrary but satisfying $C^* < C < \sup_{y \in \mathbb{R}^n} \tilde{T}_{Het}(y)$. We can then choose $y_0 \in \mathbb{R}^n$ with $\tilde{T}_{Het}(y_0) > C > 0$. In view of the definition of \tilde{T}_{Het} it follows that $y_0 \notin \tilde{\mathcal{B}}$, and hence \tilde{T}_{Het} is continuous at y_0 . We can thus find an open neighborhood $U(y_0)$ of y_0 in \mathbb{R}^n such that \tilde{T}_{Het} is larger than C on $U(y_0)$. In particular, $P_{\mu_0, \Sigma}(\tilde{T}_{Het} \geq C) \geq P_{\mu_0, \Sigma}(U(y_0)) > 0$ for every $\mu_0 \in \mathfrak{M}_0$ and every $\Sigma \in \mathfrak{C}_{Het}$. This establishes $\alpha^* > 0$. Choose $\delta > 0$ such that $\delta \leq \alpha$ and $\delta < \alpha^*$. Then the size of the rejection region $\{\tilde{T}_{Het} \geq C_\diamond(\delta)\}$ is exactly equal to δ by Part (c) of Theorem 5.1. Consequently, $\{\tilde{T}_{Het} \geq C_\diamond(\delta)\}$ is not a $\lambda_{\mathbb{R}^n}$ -null set. By construction, $C_\diamond(\alpha) \leq C_\diamond(\delta)$ holds, and hence $\{\tilde{T}_{Het} \geq C_\diamond(\alpha)\}$ contains $\{\tilde{T}_{Het} \geq C_\diamond(\delta)\}$, which completes the proof. ■

⁵⁶This argument is actually superfluous since \tilde{T}_{uc} is bounded as noted in Section 5.1.1.

Proof of Theorem A.5: We first prove Part (b). We wish to apply Part A of Proposition 5.12 of Pötscher and Preinerstorfer (2018) with $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$, $T = \tilde{T}_{Het}$, $\mathcal{L} = \mathfrak{M}_0^{lin}$, and $\mathcal{V} = \{0\}$. First, note that $\dim(\mathfrak{M}_0^{lin}) = k - q < n$. Second, under Assumption 2, \tilde{T}_{Het} is clearly Borel-measurable and is continuous on the complement of $\tilde{\mathbf{B}}$, where $\tilde{\mathbf{B}}$ is a closed $\lambda_{\mathbb{R}^n}$ -null set (see Lemma 3.3 and the paragraph following this lemma). Because of Remark 3.5, we hence see that the general assumptions on $T = \tilde{T}_{Het}$, on $N^\dagger = \tilde{\mathbf{B}}$, on $\mathcal{L} = \mathfrak{M}_0^{lin}$, as well as on $\mathcal{V} = \{0\}$ in Proposition 5.12 of Pötscher and Preinerstorfer (2018) are satisfied. Next, observe that the validity of condition (27) clearly does not depend on the choice of $\mu_0 \in \mathfrak{M}_0$ since $\tilde{\mathbf{B}} + \mathfrak{M}_0^{lin} = \tilde{\mathbf{B}}$ as shown in Lemma 3.3. For the same reason condition (27) can equivalently be written as

$$\mu_0 + \text{span} \left(\left\{ \Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n) : i \in (n_{j-1}^+, n_j^+] \right\} \right) \not\subseteq \tilde{\mathbf{B}}$$

for every $j = 1, \dots, m$, such that $(n_{j-1}^+, n_j^+] \cap I_1(\mathfrak{M}_0^{lin}) \neq \emptyset$, since $\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n)$ and $e_i(n)$ differ only by an element of \mathfrak{M}_0^{lin} . In view of Proposition B.2 in Appendix B, this implies that $\mu_0 + \mathcal{S}$ for any $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$ is not contained in $\tilde{\mathbf{B}}$, and thus not in N^\dagger . Since $\mu_0 + \mathcal{S}$ is an affine space and $N^\dagger = \tilde{\mathbf{B}}$ is a finite union of proper affine spaces under Assumption 2 as discussed in Lemma 3.3, we may conclude (cf. Corollary 5.6 in Pötscher and Preinerstorfer (2018) and its proof) that $\lambda_{\mu_0 + \mathcal{S}}(N^\dagger) = 0$ for every $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$ and every $\mu_0 \in \mathfrak{M}_0$. This completes the verification of the assumptions of Proposition 5.12 in Pötscher and Preinerstorfer (2018) that are not specific to Part A (or Part B) of this proposition. We next verify the assumptions specific to Part A of this proposition: Assumption (a) is satisfied (even for every $C \in \mathbb{R}$) as a consequence of Part (d) of Lemma D.1 (note that we have assumed that \tilde{T}_{Het} is not constant on $\mathbb{R}^n \setminus \tilde{\mathbf{B}}$). And Assumption (b) in Part A follows from Part (c) of Lemma D.1. Part A of Proposition 5.12 of Pötscher and Preinerstorfer (2018) now immediately delivers claim (26), since $C^* < \infty$ as noted in that proposition. That C^* and α^* do not depend on the choice of $\mu_0 \in \mathfrak{M}_0$ is an immediate consequence of $G(\mathfrak{M}_0)$ -invariance of \tilde{T}_{Het} . Also note that α^* as defined in the theorem coincides with α^* as defined in Proposition 5.12 of Pötscher and Preinerstorfer (2018) in view of $G(\mathfrak{M}_0)$ -invariance of \tilde{T}_{Het} . In case $\alpha < \alpha^*$, the remaining claim in Part (b) of the theorem, namely that equality can be achieved in (23), follows from the definition of C^* in Lemma 5.11 of Pötscher and Preinerstorfer (2018) and from Part A.2 of Proposition 5.12 of Pötscher and Preinerstorfer (2018) (and the observation immediately following that proposition allowing one to drop the suprema w.r.t. μ_0 and σ^2 , and to set $\sigma^2 = 1$); in case $\alpha = \alpha^* < 1$, it follows from Remarks 5.13(i),(ii) in Pötscher and Preinerstorfer (2018) using Part (d) of Lemma D.1. [In case $\alpha^* = 0$, there is nothing to prove.]

The proof of Part (a) proceeds similarly, but with some differences: Noting that \tilde{T}_{uc} is clearly Borel-measurable and is continuous on the complement of \mathfrak{M}_0 , where \mathfrak{M}_0 is a closed $\lambda_{\mathbb{R}^n}$ -null set, and using Remark 3.5, we now see that the general assumptions on $T = \tilde{T}_{uc}$, on $N^\dagger = \mathfrak{M}_0$, on $\mathcal{L} = \mathfrak{M}_0^{lin}$, as well as on $\mathcal{V} = \{0\}$ in Proposition 5.12 of Pötscher and Preinerstorfer (2018) are satisfied (again with $\mathfrak{C} = \mathfrak{C}_{(n_1, \dots, n_m)}$). Let now $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$. In view

of Proposition B.2 in Appendix B, \mathcal{S} must then contain an element of the form $\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n)$ for some $i \in I_1(\mathfrak{M}_0^{lin})$. Observe that $\Pi_{(\mathfrak{M}_0^{lin})^\perp} e_i(n) \notin \mathfrak{M}_0^{lin}$ must hold, since otherwise we would have $e_i(n) \in \mathfrak{M}_0^{lin}$, contradicting $i \in I_1(\mathfrak{M}_0^{lin})$. It follows that $\mathcal{S} \not\subseteq \mathfrak{M}_0^{lin}$, and thus $\mu_0 + \mathcal{S} \not\subseteq \mathfrak{M}_0$ for every $\mu_0 \in \mathfrak{M}_0$. Since $\mu_0 + \mathcal{S}$ is an affine space and $N^\dagger = \mathfrak{M}$ is a proper affine space we may conclude (cf. Corollary 5.6 in Pötscher and Preinerstorfer (2018) and its proof) that $\lambda_{\mu_0 + \mathcal{S}}(N^\dagger) = 0$ for every $\mathcal{S} \in \mathbb{J}(\mathfrak{M}_0^{lin}, \mathfrak{C}_{(n_1, \dots, n_m)})$ and every $\mu_0 \in \mathfrak{M}_0$. We have thus now completed the verification of the assumptions of Proposition 5.12 of Pötscher and Preinerstorfer (2018) that are not specific to Part A (or Part B) of this proposition. We next verify the assumptions specific to Part A of this proposition: Verification of Assumptions (a) and (b) in Part A of Proposition 5.12 of Pötscher and Preinerstorfer (2018) proceeds similar as before except for now using Parts (b) and (a) of Lemma D.1. Part A of Proposition 5.12 of Pötscher and Preinerstorfer (2018) now immediately delivers claim (25), again since $C^* < \infty$ as noted in that proposition.⁵⁷ Again, $G(\mathfrak{M}_0)$ -invariance of \tilde{T}_{uc} implies that C^* and α^* do not depend on the choice of $\mu_0 \in \mathfrak{M}_0$, and that α^* as defined in the theorem coincides with α^* as defined in Proposition 5.12 of Pötscher and Preinerstorfer (2018). The remaining claim in Part (a) is proved completely analogous as the corresponding claim in Part (b) except for now using Part (b) of Lemma D.1.

We finally prove Part (c): The first claim follows from Remark 5.10 in Pötscher and Preinerstorfer (2018) and Lemma D.1. The second claim is obvious. ■

E Appendix: Algorithms

In this appendix, we discuss algorithms for determining rejection probabilities and the size of a test based on one of the test statistics T_{Het} , T_{uc} , \tilde{T}_{Het} , or \tilde{T}_{uc} together with a given candidate critical value, as well as size-controlling critical values. We discuss these algorithms under the Gaussianity assumption made in Section 2, but recall from Section 6.1 that the algorithms as given here can also be used to calculate *null* rejection probabilities, size, and size-controlling critical values in the elliptically symmetric case *without any changes*. Furthermore, we restrict ourselves to the heteroskedasticity model \mathfrak{C}_{Het} ; adapting the algorithms to subsets \mathfrak{C} of \mathfrak{C}_{Het} is rather straightforward (basically one has to appropriately constrain the optimization routines involved, appropriately redefine some of the quantities like C_{low} , and refer to the size-control conditions pertinent for the given heteroskedasticity model \mathfrak{C}).

E.1 Computing rejection probabilities

Suppose that a $G(\mathfrak{M}_0)$ -invariant test statistic $T : \mathbb{R}^n \rightarrow \mathbb{R}$ has the following property: for some (and hence any) $\mu_0 \in \mathfrak{M}_0$ and a critical value $C \in \mathbb{R}$, there exists a symmetric $n \times n$ matrix A_C ,

⁵⁷This argument is actually superfluous since \tilde{T}_{uc} is bounded as noted in Section 5.1.1.

such that

$$T(\mu_0 + z) \geq C \Leftrightarrow z' A_C z \geq 0 \text{ holds for } \lambda_{\mathbb{R}^n}\text{-almost every } z \in \mathbb{R}^n. \quad (31)$$

If this property is satisfied, then for all choices of $\Sigma \in \mathfrak{C}_{Het}$, $\mu_0 \in \mathfrak{M}_0$, $\mu \in \mathfrak{M}$, and $\sigma^2 \in (0, \infty)$, and setting $\nu := \sigma^{-1} \Sigma^{-1/2}(\mu - \mu_0)$, we may write

$$P_{\mu, \sigma^2 \Sigma}(\{z \in \mathbb{R}^n : T(z) \geq C\}) = P_{\nu, I_n}(\{\zeta \in \mathbb{R}^n : \zeta' \Sigma^{1/2} A_C \Sigma^{1/2} \zeta \geq 0\}); \quad (32)$$

in case $\mu \in \mathfrak{M}_0$, we may set $\mu_0 = \mu$ to further simplify the right-hand-side in (32) to

$$P_{0, I_n}(\{\zeta \in \mathbb{R}^n : \zeta' \Sigma^{1/2} A_C \Sigma^{1/2} \zeta \geq 0\}). \quad (33)$$

The probability that a Gaussian quadratic form is not less than 0 (such as (32) or (33)) can numerically be determined by standard algorithms such as Davies (1980). Relation (31) can thus be exploited for efficiently computing rejection probabilities (for a given critical value), and thus plays an instrumental rôle in determining the size of a test, size-controlling critical values, or the power function of a test.

For the important case $q = 1$ we now show that the above approach can indeed be used. It follows from the subsequent lemma that for any critical value C the property in (31) holds for the following test statistics: (i) T_{Het} provided Assumption 1 holds; (ii) T_{uc} ; (iii) \tilde{T}_{Het} provided Assumption 2 holds; (iv) \tilde{T}_{uc} . Recall from Lemmata 3.1 and 3.3 that under Assumption 1 (Assumption 2, respectively), the set B (\tilde{B} , respectively) is a $\lambda_{\mathbb{R}^n}$ -null set. Note that v defined in the lemma satisfies $v \neq 0$.

Lemma E.1. *Suppose $q = 1$. Let $v = v_{R,X} := X(X'X)^{-1}R'$. Then, for every $C \in \mathbb{R}$ and every $\mu_0 \in \mathfrak{M}_0$, we have:*

(a) *If $\mu_0 + z \notin B$, then $T_{Het}(\mu_0 + z) \geq C$ ($\leq C$) is equivalent to $z' A_{Het,C} z \geq 0$ (≤ 0), where*

$$A_{Het,C} := vv' - C \Pi_{\text{span}(X)^\perp} \text{diag}(v_1^2 d_1, \dots, v_n^2 d_n) \Pi_{\text{span}(X)^\perp}. \quad (34)$$

(b) *If $\mu_0 + z \notin \text{span}(X)$, then $T_{uc}(\mu_0 + z) \geq C$ ($\leq C$) is equivalent to $z' A_{uc,C} z \geq 0$ (≤ 0), where*

$$A_{uc,C} := vv' - C \frac{v'v}{n-k} \Pi_{\text{span}(X)^\perp}. \quad (35)$$

(c) *If $\mu_0 + z \notin \tilde{B}$, then $\tilde{T}_{Het}(\mu_0 + z) \geq C$ ($\leq C$) is equivalent to $z' \tilde{A}_{Het,C} z \geq 0$ (≤ 0), where*

$$\tilde{A}_{Het,C} := vv' - C \Pi_{(\mathfrak{M}_0^{lin})^\perp} \text{diag}(v_1^2 \tilde{d}_1, \dots, v_n^2 \tilde{d}_n) \Pi_{(\mathfrak{M}_0^{lin})^\perp}. \quad (36)$$

(d) *If $\mu_0 + z \notin \mathfrak{M}_0$, then $\tilde{T}_{uc}(\mu_0 + z) \geq C$ ($\leq C$) is equivalent to $z' \tilde{A}_{uc,C} z \geq 0$ (≤ 0), where*

$$\tilde{A}_{uc,C} := vv' - C \frac{v'v}{n-(k-1)} \Pi_{(\mathfrak{M}_0^{lin})^\perp}. \quad (37)$$

Proof: We first observe that there is nothing to prove in Part (a) (Part (c), respectively) if Assumption 1 (Assumption 2, respectively) is violated, since then $B = \mathbb{R}^n$ ($\tilde{B} = \mathbb{R}^n$, respectively) by Lemma 3.1 (Lemma 3.3, respectively). In the following we hence may assume for Part (a) (Part (c), respectively) that Assumption 1 (Assumption 2, respectively) hold, in which case B (\tilde{B} , respectively) is a $\lambda_{\mathbb{R}^n}$ -null set. The expressions in (34), (35), (36), and (37) now follow directly from the definitions of the test statistics since $q = 1$, recalling in particular that $\hat{u}(\mu_0 + z) = \Pi_{\text{span}(X)^\perp}(\mu_0 + z) = \Pi_{\text{span}(X)^\perp}z$, and $\tilde{u}(\mu_0 + z) = \Pi_{(\mathfrak{M}_0^{\text{lin}})^\perp}((\mu_0 + z) - \mu_0) = \Pi_{(\mathfrak{M}_0^{\text{lin}})^\perp}z$, and noting that for $q = 1$

$$\begin{aligned} R\hat{\beta}(\mu_0 + z) &= r + v'z, \\ \hat{\Omega}_{Het}(\mu_0 + z) &= z'\Pi_{\text{span}(X)^\perp} \text{diag}(v_1^2 d_1, \dots, d_n v_n^2) \Pi_{\text{span}(X)^\perp} z, \\ \tilde{\Omega}_{Het}(\mu_0 + z) &= z'\Pi_{(\mathfrak{M}_0^{\text{lin}})^\perp} \text{diag}(v_1^2 \tilde{d}_1, \dots, \tilde{d}_n v_n^2) \Pi_{(\mathfrak{M}_0^{\text{lin}})^\perp} z, \\ \hat{\sigma}^2(\mu_0 + z) &= \frac{z'\Pi_{\text{span}(X)^\perp} z}{n - k}, \quad \tilde{\sigma}^2(z) = \frac{z'\Pi_{(\mathfrak{M}_0^{\text{lin}})^\perp} z}{n - (k - 1)} \end{aligned}$$

hold. ■

Remark E.2. The algorithm in Davies (1980) applied to (32) requires that the matrix A_C is not the zero matrix.

(i) It is easy to see that $A_{Het,C}$, $A_{uc,C}$, and $\tilde{A}_{uc,C}$ are never equal to the zero matrix: Note that $v'A_{Het,C}v = (v'v)^2 > 0$, since $v \in \text{span}(X)$ and $v \neq 0$. The same argument applies to $A_{uc,C}$. Furthermore, for $C = 0$ the matrix $\tilde{A}_{uc,C}$ is obviously not the zero matrix; for $C \neq 0$ let $w \in (\mathfrak{M}_0^{\text{lin}})^\perp$, $w \neq 0$, w orthogonal to v , then $w'\tilde{A}_{uc,C}w = -w'wCv'v/(n - (k - 1)) \neq 0$ (note that such a w exists, since $v \in (\mathfrak{M}_0^{\text{lin}})^\perp$ and $\dim((\mathfrak{M}_0^{\text{lin}})^\perp) = n - (k - q) > n - k \geq 1$ hold).

(ii) For $\tilde{A}_{Het,C}$ we have the following: Since $v \in (\mathfrak{M}_0^{\text{lin}})^\perp$ holds, $v'\tilde{A}_{Het,C}v = (v'v)^2 - Cv' \text{diag}(v_1^2 \tilde{d}_1, \dots, v_n^2 \tilde{d}_n)v$, which is zero only for $C = C_0$ where $C_0 = \sum_{i=1}^n v_i^2 / \sum_{i=1}^n v_i^4 \tilde{d}_i$ (note that the ratio is well-defined since all the \tilde{d}_i are positive and since $v \neq 0$). Hence, $\tilde{A}_{Het,C}$ is not the zero matrix, except possibly for $C = C_0$. We now show that – in case Assumption 2 is satisfied – $\tilde{A}_{Het,C_0} = 0$ is equivalent to $\tilde{T}_{Het}(y)$ being constant for $y \in \mathbb{R}^n \setminus \tilde{B}$: Suppose $\tilde{A}_{Het,C_0} = 0$. Since $C_0 > 0$, we obtain $\Pi_{(\mathfrak{M}_0^{\text{lin}})^\perp} \text{diag}(v_1^2 \tilde{d}_1, \dots, v_n^2 \tilde{d}_n) \Pi_{(\mathfrak{M}_0^{\text{lin}})^\perp} = vv'/C_0$ and thus $\tilde{A}_{Het,C} = vv'(1 - C/C_0)$. Fix $\mu_0 \in \mathfrak{M}_0$ arbitrary. For every $C > C_0$ we have $z'\tilde{A}_{Het,C}z \leq 0$ for every z , and hence for every z with $\mu_0 + z \notin \tilde{B}$ (note that $\mathbb{R}^n \setminus \tilde{B}$ is nonempty under Assumption 2). By Lemma E.1 we can conclude that $\tilde{T}_{Het}(\mu_0 + z) \leq C$ for every $\mu_0 + z \notin \tilde{B}$. By the same token, we obtain that $\tilde{T}_{Het}(\mu_0 + z) \geq C$ for every $\mu_0 + z \notin \tilde{B}$ when $C < C_0$ holds. We conclude that $\tilde{T}_{Het}(\mu_0 + z) = C_0$ for every $\mu_0 + z \notin \tilde{B}$, i.e., $\tilde{T}_{Het}(y) = C_0$ for every $y \in \mathbb{R}^n \setminus \tilde{B}$. To prove the converse, assume $\tilde{T}_{Het}(y) = C_1$ for every $y \notin \tilde{B}$. Fix $\mu_0 \in \mathfrak{M}_0$ arbitrary. Then $\tilde{T}_{Het}(\mu_0 + z) = C_1$ for every z with $\mu_0 + z \notin \tilde{B}$. By Lemma E.1 we get $z'\tilde{A}_{Het,C}z \geq 0$ (≤ 0 , respectively) for $C \leq C_1$ ($C \geq C_1$, respectively) for every $z \notin \tilde{B} - \mu_0$. Under Assumption 2 the set $\tilde{B} - \mu_0$ is a $\lambda_{\mathbb{R}^n}$ -null set, hence its complement is dense in \mathbb{R}^n . By continuity of the quadratic forms, we get $z'\tilde{A}_{Het,C}z \geq 0$ (≤ 0 , respectively) for $C \leq C_1$ ($C \geq C_1$, respectively) for all $z \in \mathbb{R}^n$. We thus

obtain $z' \tilde{A}_{Het, C_1} z = 0$ for every $z \in \mathbb{R}^n$. Since \tilde{A}_{Het, C_1} is symmetric, $\tilde{A}_{Het, C_1} = 0$ follows and $C_1 = C_0$ must hold.

(iii) Before applying the algorithm in Davies (1980) to (32) with $T = T_{Het}$ and $A_C = A_{Het, C}$ we first check that Assumption 1 holds since otherwise Part (a) of the preceding lemma does not apply. In case of $T = \tilde{T}_{Het}$ and $A_C = \tilde{A}_{Het, C}$ we check that Assumption 2 holds for similar reasons; and, in case this assumption is satisfied, we then always also compute C_0 and check numerically that \tilde{A}_{Het, C_0} (and hence any $\tilde{A}_{Het, C}$) is not the zero matrix.

In case $q > 1$, the algorithm in Davies (1980) could also be used to compute rejection probabilities for the tests based on T_{uc} and \tilde{T}_{uc} as is easy to see. Since this is not so for T_{Het} and \tilde{T}_{Het} , we do not proceed in this way for reasons of comparability. In case $q > 1$ we thus compute the required rejection probabilities by Monte Carlo.

E.2 Determining the size of a test

For simplicity throughout this subsection T denotes any one of the test statistics UC , $HC0-HC4$, UCR , $HC0R-HC4R$. In case of $HC0-HC4$ we assume in our discussion that the design matrix X and R are such that Assumption 1 is satisfied, and in case of $HC0R-HC4R$ we assume that Assumption 2 holds and that the test statistic is not constant on $\mathbb{R}^n \setminus \tilde{\mathcal{B}}$.⁵⁸ These conditions should be checked either theoretically or numerically before using the algorithms described below. Such numerical checks are implemented in the R-package **hrt** (Preinerstorfer (2021)) realizing these algorithms.

We now discuss algorithms for determining the size (over \mathfrak{C}_{Het}) of the test that rejects if $T \geq C$ for a given critical value $C > 0$ (note that any $C \leq 0$ leads to a trivial test that always rejects). By $G(\mathfrak{M}_0)$ -invariance of T , for any given $\mu_0 \in \mathfrak{M}_0$, the size of this test simplifies to

$$\sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \Sigma}(T \geq C), \quad (38)$$

which is what the algorithms described below compute numerically.

Before trying to determine the size numerically, it is advisable to check whether C is not less than the pertinent lower bound C^* for size-controlling critical values obtained in our theoretical results in Propositions 4.5 and 5.4 (and the attending footnotes), since otherwise one already knows that the size of the test is equal to 1, and hence there is no need to run the algorithm. The implementations of the algorithms in the R-package **hrt** (Preinerstorfer (2021)) have an option that provides such a check and outputs 1 if the check fails without running the algorithm.

Of course, the design matrix X , the restriction (R, r) , and the particular choice of test statistic from the above list, are inputs to all the algorithms that are discussed in this and the subsequent section E.3, but we do not show these inputs explicitly in the descriptions of the algorithms given further down.

⁵⁸This rules out trivial cases only.

E.2.1 Case $q = 1$

In the important special case $q = 1$ we can use (33) and Lemma E.1 to compute the rejection probabilities $P_{\mu_0, \Sigma}(T \geq C)$ appearing in (38) efficiently via, e.g., Davies (1980) (referred to as DA in what follows). A generic algorithm based on this observation is summarized in Algorithm 1.

Algorithm 1 Computing the size in case $q = 1$.

```

1: Input A real number  $C > 0$  and positive integers  $M_0 \geq M_1 \geq M_2$ .
2: Stage 0: Initial value search
3: for  $j = 1$  to  $j = M_0$  do
4:   Choose a candidate  $\Sigma_j \in \mathfrak{C}_{Het}$ .
5:   Obtain  $\tilde{p}_j := P_{\mu_0, \Sigma_j}(T \geq C)$  using DA.
6: end for
7: Rank the candidates  $\Sigma_j$  according to the value (from largest to smallest) of the corresponding
   quantities  $\tilde{p}_j$  to obtain  $\Sigma_{1:M_0}, \dots, \Sigma_{M_1:M_0}$ , the initial values for the next stage.
8: Stage 1: Coarse localized optimizations
9: for  $j = 1$  to  $j = M_1$  do
10:  Obtain  $\Sigma_j^*$  by running a numerical algorithm for the optimization problem (38) initialized
    at  $\Sigma_{j:M_0}$  and obtain  $\bar{p}_{j, \Sigma_j^*} := P_{\mu_0, \Sigma_j^*}(T \geq C)$  (using DA to evaluate probabilities).
11: end for
12: Rank the obtained matrices  $\Sigma_j^*$  according to the value (from largest to smallest) of the
    corresponding  $\bar{p}_{j, \Sigma_j^*}$  to obtain  $\Sigma_{1:M_1}^*, \dots, \Sigma_{M_2:M_1}^*$ , the initial values for the next stage.
13: Stage 2: Refined localized optimization
14: for  $j = 1$  to  $j = M_2$  do
15:  Obtain  $\Sigma_j^{**}$  by running a (refined) numerical algorithm for the optimization problem (38)
    initialized at  $\Sigma_{j:M_1}^*$  and obtain  $\bar{\bar{p}}_{j, \Sigma_j^{**}} := P_{\mu_0, \Sigma_j^{**}}(T \geq C)$  (using DA to evaluate probabilities).
16: end for
17: Return  $\max_{j=1, \dots, M_2} \bar{\bar{p}}_{j, \Sigma_j^{**}}$ 

```

Remark E.3. The initial values Σ_j in Stage 0 of Algorithm 1 can, for example, be obtained randomly (e.g., by sampling the diagonal elements of Σ_j from a uniform distribution on the unit simplex in \mathbb{R}^n). Such random choices may then be supplemented by “special” elements of \mathfrak{C}_{Het} , e.g., matrices that are close to $e_i(n)e_i(n)'$, $i = 1, \dots, n$, or the matrix $n^{-1}I_n$, or a matrix Σ that maximizes the expectation of the quadratic form $y \mapsto y'\Sigma^{1/2}A_C\Sigma^{1/2}y$ under P_{0, I_n} (where A_C is obtained via Lemma E.1, cf. also the discussion preceding that lemma), the latter choice being motivated by (33). For the particular choice of initial values used in the R-package **hrt** and in our numerical calculations see Preinerstorfer (2021) and Appendix F.

Remark E.4. If Algorithm 1 is to be applied to a relatively large critical value C (say C larger than 5 times the $(1 - \alpha)$ -quantile of the cdf of $P_{0, I_N} \circ T$), then one may run Algorithm 1 on a smaller critical value first (e.g., the just mentioned quantile), and use the covariance matrix realizing the maximal rejection probability for this smaller critical value (in line 17 of Algorithm 1) as an additional initial value when running Algorithm 1 for determining the size corresponding

to the originally given C . This can help to ameliorate numerical difficulties due to the rejection probabilities being close to zero over large portions of \mathfrak{C}_{Het} . The just described procedure is available as an option in the R-package **hrt**.

Remark E.5. The concrete choice of the numerical optimization algorithm used in Stages 1 and 2 of Algorithm 1 is left unspecified here, but may, for example, be a constrained Nelder and Mead (1965) algorithm (as provided in R’s “constrOptim” function), where in Stage 2 the parameters in this algorithm (and in principle also in DA) should be chosen to guarantee a higher accuracy. For the particular choice of optimization routines used in the R-package **hrt** and in our numerical calculations see Preinerstorfer (2021) and Appendix F.

Remarks E.3 and E.5 also apply to other algorithms introduced further down, and will not be repeated.

E.2.2 General case

An algorithm that is similar to Algorithm 1, but uses Monte-Carlo simulation instead of DA to compute the rejection probabilities $P_{\mu_0, \Sigma}(T \geq C)$ is discussed in Algorithm 2; this algorithm is a modification of Algorithm 2 in Pötscher and Preinerstorfer (2018).⁵⁹ In Algorithm 2 the number of replications used in the Monte-Carlo simulations (and thus their accuracy but also their runtime) is increased in each stage, leading to an improved accuracy in the rejection probabilities computed. While this algorithm is also applicable in case $q = 1$, Algorithm 1 is to be preferred (and is automatically applied by the R-package **hrt** in this case), as it is based on a preferable way of computing the rejection probabilities.

E.3 Determining size-controlling critical values

Again, in this subsection T denotes any one of the test statistics UC , $HC0-HC4$, UCR , $HC0R-HC4R$. In case of $HC0-HC4$ we assume in our discussion that the design matrix X and R are such that Assumption 1 is satisfied, and in case of $HC0R-HC4R$ we assume that Assumption 2 holds and that the test statistic is not constant on $\mathbb{R}^n \setminus \tilde{\mathcal{B}}$.⁶⁰ Furthermore, we assume that size-controlling critical values exist. These conditions should be checked either theoretically or numerically before using the algorithms described below. The last mentioned existence can be guaranteed by checking (theoretically or numerically) the respective sufficient conditions for size control in Theorems 4.1 and 5.1.⁶¹ We note that the implementations of the algorithms presented below in the R-package **hrt** (Preinerstorfer (2021)) include such numerical checks.

⁵⁹This algorithm involves evaluating the test statistic T . Since the definition of T depends on invertibility of a covariance matrix estimator, an invertibility check is required. We use the same invertibility check as discussed in the second paragraph in Appendix E.3 of Pötscher and Preinerstorfer (2020), with a tolerance parameter that can be specified by the user.

⁶⁰This rules out trivial cases only.

⁶¹In case the respective sufficient conditions are violated, but size-controlling critical values nevertheless exist (as, e.g., in Example 4.5), the algorithm still works.

Algorithm 2 Computing the size for general q .

- 1: **Input** A real number $C > 0$ and positive integers $M_0 \geq M_1 \geq M_2$, $N_0 \leq N_1 \leq N_2$.
 - 2: **Stage 0: Initial value search**
 - 3: **for** $j = 1$ **to** $j = M_0$ **do**
 - 4: Generate a pseudorandom sample Z_1, \dots, Z_{N_0} from P_{0, I_n} .
 - 5: Obtain a candidate $\Sigma_j \in \mathfrak{C}_{Het}$.
 - 6: Compute $\tilde{p}_j = N_0^{-1} \sum_{i=1}^{N_0} \mathbf{1}_{[C, \infty)}(T(\mu_0 + \Sigma_j^{1/2} Z_i))$.
 - 7: **end for**
 - 8: Rank the candidates Σ_j according to the value (from largest to smallest) of the corresponding quantities \tilde{p}_j to obtain $\Sigma_{1:M_0}, \dots, \Sigma_{M_1:M_0}$, the initial values for the next stage.
 - 9: **Stage 1: Coarse localized optimizations**
 - 10: **for** $j = 1$ **to** $j = M_1$ **do**
 - 11: Generate a pseudorandom sample Z_1, \dots, Z_{N_1} from P_{0, I_n} .
 - 12: Define $\bar{p}_{j, \Sigma} = N_1^{-1} \sum_{i=1}^{N_1} \mathbf{1}_{[C, \infty)}(T(\mu_0 + \Sigma^{1/2} Z_i))$ for $\Sigma \in \mathfrak{C}_{Het}$.
 - 13: Obtain Σ_j^* by running a numerical optimization algorithm for the problem $\sup_{\Sigma \in \mathfrak{C}_{Het}} \bar{p}_{j, \Sigma}$ initialized at $\Sigma_{j:M_0}$.
 - 14: **end for**
 - 15: Rank the obtained numbers Σ_j^* according to the value (from largest to smallest) of the corresponding \bar{p}_{j, Σ_j^*} to obtain $\Sigma_{1:M_1}^*, \dots, \Sigma_{M_2:M_1}^*$, the initial values for the next stage.
 - 16: **Stage 2: Refined localized optimization**
 - 17: **for** $j = 1$ **to** $j = M_2$ **do**
 - 18: Generate a pseudorandom sample Z_1, \dots, Z_{N_2} from P_{0, I_n} .
 - 19: Define $\bar{\bar{p}}_{j, \Sigma} = N_2^{-1} \sum_{i=1}^{N_2} \mathbf{1}_{[C, \infty)}(T(\mu_0 + \Sigma^{1/2} Z_i))$ for $\Sigma \in \mathfrak{C}_{Het}$.
 - 20: Obtain Σ_j^{**} by running a numerical optimization algorithm for the problem $\sup_{\Sigma \in \mathfrak{C}_{Het}} \bar{\bar{p}}_{j, \Sigma}$ initialized at $\Sigma_{j:M_1}^*$.
 - 21: **end for**
 - 22: **Return** $\max_{j=1, \dots, M_2} \bar{\bar{p}}_{j, \Sigma_j^{**}}$.
-

We now proceed to discussing several algorithms for determining the smallest critical value $C_\diamond(\alpha)$ such that the size of the test, which rejects if $T \geq C_\diamond(\alpha)$, does not exceed α ($0 < \alpha < 1$).⁶² [In fact, for $C_\diamond(\alpha)$ the size then equals α *provided* a critical value that results in size *equal* to α actually exists.] Note that $C_\diamond(\alpha) > 0$ must hold, in view of Remarks 4.4 and 5.3 since $\alpha < 1$. By $G(\mathfrak{M}_0)$ -invariance, for some fixed $\mu_0 \in \mathfrak{M}_0$, the algorithms numerically compute the smallest critical value that satisfies

$$\sup_{\Sigma \in \mathfrak{C}_{Het}} P_{\mu_0, \Sigma}(T \geq C) \leq \alpha, \quad (39)$$

cf. the discussion surrounding (38). For later use we denote by F_Σ the cdf of $P_{\mu_0, \Sigma} \circ T$, which by $G(\mathfrak{M}_0)$ -invariance does not depend on the particular choice for $\mu_0 \in \mathfrak{M}_0$.

E.3.1 Computing size-controlling critical values via line search based on algorithms in Section E.2

Given an algorithm $A : (0, \infty) \rightarrow [0, 1]$ that for $C > 0$ returns the size of the test that rejects if $T \geq C$, one can use a line-search algorithm to determine the smallest critical value $C = C_\diamond(\alpha)$ satisfying $A(C) \leq \alpha$. To this end, one starts at the lower bound $C_{low} = \max(C^*, C_{hom})$, where C^* is given in the pertinent parts of Theorems 4.1 and 5.1, respectively, and C_{hom} denotes the *smallest* $1 - \alpha$ quantile of F_{I_n} , i.e., of the cdf of the test statistic under homoskedasticity. Note that then $P_{\mu_0, I_n}(T \geq C_{hom}) = \alpha$ and that $P_{\mu_0, I_n}(T \geq C) > \alpha$ for $C < C_{hom}$ (to see this note that F_{I_n} is continuous as $\{T = C\}$ is a $\lambda_{\mathbb{R}^n}$ -null set for all real C , cf. Lemma 5.16 in Pötscher and Preinerstorfer (2018) and Lemma D.1 in Appendix D). Furthermore, $C_{hom} > 0$ (since $T \geq 0$ and $\{T = 0\}$ is a $\lambda_{\mathbb{R}^n}$ -null set), and consequently $C_{low} > 0$ holds. Starting from C_{low} , one then keeps increasing the critical value “in a reasonable way” until one obtains, for the first time, a C such that $A(C) \leq \alpha$ holds. This procedure is summarized in Algorithm 3, in which the particular algorithm A used is an input to Algorithm 3. For A one may either use Algorithm 1 if $q = 1$, or Algorithm 2 for general q . Note that one may need to terminate the while-loop after a maximal number of iterations.

Remark E.6. (i) Note that a matrix Σ^{**} as required for the while-loop in Algorithm 3 can easily be obtained by implementing Algorithm 1 or 2 in such a fashion as to also return the covariance matrix for which the maximal rejection probability is attained in the respective Stage 2.

(ii) A smallest C_+ as required in line 5 of Algorithm 3 indeed exists since $\{T = C\}$ is a $\lambda_{\mathbb{R}^n}$ -null set for all real C as noted before.

(iii) For details regarding the computation of C_{low} in the R-package **hrt** see Preinerstorfer (2021) and Appendix F.2.

⁶²Such a smallest size-controlling critical value indeed exists under the assumptions of this subsection in view of Remark 5.10 and Lemma 5.16 in Pötscher and Preinerstorfer (2018) and Lemma D.1 in Appendix D. [Under the sufficient conditions for size control in the respective theorems, this can also be read off directly from these theorems.]

Algorithm 3 Numerical approximation of the smallest size-controlling critical value via a line search algorithm.

```

1: Input  $\alpha \in (0, 1)$ ,  $A$ ,  $C_{low}$ ,  $\epsilon \in [0, 1 - \alpha)$  ( $\epsilon$  a small tolerance parameter).
2:  $C \leftarrow C_{low}$ 
3: while  $A(C) > \alpha + \epsilon$  do
4:   Let  $\Sigma^{**}$  be such that  $P_{\mu_0, \Sigma^{**}}(T \geq C) \approx A(C)$ .
5:   Determine, by an upward line search initialized at  $C$ , the smallest value  $C_+$  such
     that  $P_{\mu_0, \Sigma^{**}}(T \geq C_+) \leq \alpha$ .
6:    $C \leftarrow C_+$ .
7: end while
8: return  $C$ 

```

E.3.2 Computing size-controlling critical values via quantile maximization

For completeness and comparison with Pötscher and Preinerstorfer (2018), we briefly describe an algorithm that is a modification of Algorithm 1 in Pötscher and Preinerstorfer (2018). In contrast to the algorithm discussed in the previous section, it does not make use of size-computations, but determines the smallest size-controlling critical value as

$$\sup_{\Sigma \in \mathfrak{C}_{Het}} F_{\Sigma}^{-1}(1 - \alpha) \quad (40)$$

where F_{Σ}^{-1} denotes the quantile function of the cdf F_{Σ} . That (40) indeed gives the smallest size-controlling critical value is not difficult to see keeping in mind that $P_{\mu_0, \Sigma}(T = C) = 0$ for every real C , every $\mu_0 \in \mathfrak{M}_0$, and every $\Sigma \in \mathfrak{C}_{Het}$ (in view of $\lambda_{\mathbb{R}^n}(\{T = C\}) = 0$ as noted before). The algorithm is summarized in Algorithm 4.

F Appendix: Details concerning numerical computations in Section 9

F.1 Details concerning Section 9.1

To obtain Tables 1 and 2, for each of the test statistics UC, HC0-HC4, UCR, HC0R-HC4R, we repeated the procedure summarized in Algorithm 5 below 15 times (recall that $n = 25$, and $R = (0, 1)$). Each time this algorithm returned a design matrix, the corresponding size of the rejection region $\{T \geq C_{\chi^2, 0.05}\}$ was obtained for the specific test statistic used, as well as a corresponding lower bound for the smallest size-controlling critical value. Then, we computed the maximum out of the 15 lower bounds, which (for each test statistic) is reported in Table 1. We also computed the maximum out of the 15 sizes, which (for each test statistic) is reported in Table 2. We also did the same with the critical value $C_{\chi^2, 0.05}$ replaced by the 95%-quantile of an $F_{1, n-k}$ -distribution ($n - k = 23$), the corresponding results being reported in Table 3.

In the description of Algorithm 5, the function $f(x)$ is an abbreviation for $C^* = \max\{T(\mu_0 +$

Algorithm 4 Numerical approximation of the smallest size-controlling critical value via quantiles.

```

1: Input Positive integers  $M_0 \geq M_1 \geq M_2$ ,  $N_0 \leq N_1 \leq N_2$ .
2: Stage 0: Initial value search
3: for  $j = 1$  to  $j = M_0$  do
4:   Generate a pseudorandom sample  $Z_1, \dots, Z_{N_0}$  from  $P_{0, I_n}$ .
5:   Obtain a candidate  $\Sigma_j \in \mathfrak{C}_{Het}$ .
6:   Compute  $\tilde{F}_j^{-1}(1 - \alpha)$  where  $\tilde{F}_j(x) = N_0^{-1} \sum_{i=1}^{N_0} \mathbf{1}_{(-\infty, x]}(T(\mu_0 + \Sigma_j^{1/2} Z_i))$  for  $x \in \mathbb{R}$ .
7: end for
8: Rank the candidates  $\Sigma_j$  according to the value (from largest to smallest) of the corresponding
   quantities  $\tilde{F}_j^{-1}(1 - \alpha)$  to obtain  $\Sigma_{1:M_0}, \dots, \Sigma_{M_1:M_0}$ , the initial values for the next stage.
9: Stage 1: Coarse localized optimizations
10: for  $j = 1$  to  $j = M_1$  do
11:   Generate a pseudorandom sample  $Z_1, \dots, Z_{N_1}$  from  $P_{0, I_n}$ .
12:   Define  $\bar{F}_{j, \Sigma}(x) = N_1^{-1} \sum_{i=1}^{N_1} \mathbf{1}_{(-\infty, x]}(T(\mu_0 + \Sigma^{1/2} Z_i))$  for  $x \in \mathbb{R}$  and  $\Sigma \in \mathfrak{C}_{Het}$ .
13:   Obtain  $\Sigma_j^*$  by running a numerical optimization algorithm for the problem
      $\sup_{\Sigma \in \mathfrak{C}_{Het}} \bar{F}_{j, \Sigma}^{-1}(1 - \alpha)$  initialized at  $\Sigma_{j:M_0}$ .
14: end for
15: Rank the obtained  $\Sigma_j^*$  according to the value (from largest to smallest) of the corresponding
      $\bar{F}_{j, \Sigma_j^*}^{-1}(1 - \alpha)$  to obtain  $\Sigma_{1:M_1}^*, \dots, \Sigma_{M_2:M_1}^*$ , the initial values for the next stage.
16: Stage 2: Refined localized optimization
17: for  $j = 1$  to  $j = M_2$  do
18:   Generate a pseudorandom sample  $Z_1, \dots, Z_{N_2}$  from  $P_{0, I_n}$ .
19:   Define  $\bar{\bar{F}}_{j, \Sigma}(x) = N_2^{-1} \sum_{i=1}^{N_2} \mathbf{1}_{(-\infty, x]}(T(\mu_0 + \Sigma^{1/2} Z_i))$  for  $x \in \mathbb{R}$  and  $\Sigma \in \mathfrak{C}_{Het}$ .
20:   Obtain  $\Sigma_j^{**}$  by running a numerical optimization algorithm for the problem
      $\sup_{\Sigma \in \mathfrak{C}_{Het}} \bar{\bar{F}}_{j, \Sigma}^{-1}(1 - \alpha)$  initialized at  $\Sigma_{j:M_1}^*$ .
21: end for
22: Return  $\max_{j=1, \dots, M_2} \bar{\bar{F}}_{j, \Sigma_j^{**}}^{-1}(1 - \alpha)$ .

```

$e_i(n) : i \in I_1(\mathfrak{M}_0^{lin})\}$, the lower bound for the size-controlling critical values (cf. Propositions 4.5, 5.4, and the attending footnotes), with the $n \times 2$ design matrix X given by an intercept ι , say, as the first column and a regressor x as the second one. Note that computing C^* necessitates the evaluation of the test statistic on a finite set of elements of \mathbb{R}^n , and then determining the maximum among the values obtained.⁶³ Concerning the evaluation of test statistics, the definition of which depends on the invertibility of a covariance matrix estimator, we used the same invertibility check as discussed in the second paragraph in Appendix E.3 of Pötscher and Preinerstorfer (2020) with a tolerance parameter of 10^{-8} . For $R = (0, 1)$ and for each matrix X returned by Algorithm 5 all relevant assumptions (i.e., the assumptions in the pertinent parts of Theorems 4.1 and 5.1, respectively) have been checked numerically.

Algorithm 5 Search procedure used for generating Tables 1 and 2.

```

1: Initialize  $x \leftarrow 0 \in \mathbb{R}^n$ .
2: for  $i = 1$  to  $i = 5$  do
3:   Generate an  $n$ -dimensional pseudo-random vector  $z$  of independent coordinates each from
   a log-standard normal distribution.
4:   Run a Nelder and Mead (1965) algorithm initialized at  $z$  to maximize  $f$  over  $\mathbb{R}^n$  (with
   a maximal number of iterations of 50, and otherwise the default parameters in R's "optim"
   function) to obtain  $z^*$ , say.
5:   if  $i = 1$ , or  $i \geq 2$  and  $f(z^*) > f(x)$  then
6:      $x \leftarrow z^*$ .
7:   end if
8:   if  $f(x) > 4$  then
9:     Go to line 12.
10:  end if
11: end for
12: Use Algorithm 1 to determine the size of the test for the test statistic under consideration for
   the design matrix  $(\iota, x)$  and based on either of the following two critical values: (i)  $C_{\chi^2, 0.05}$ 
   and (ii) the 95% quantile of an  $F_{1, n-k}$  distribution.
13: return  $x$ ,  $f(x)$ , and the two sizes determined in the previous step.
```

Algorithm 5 uses Algorithm 1 in determining the size of a given test. We made the following choices concerning the parameters required in Algorithm 1 (and used default settings if not mentioned otherwise):

1. The candidates in Stage 0 of Algorithm 1 were determined by combining the suggestions in Remarks E.3 and E.4. That is, denoting $M_p = 200\,000$, we combined: (i) sampling $M_p/4 - 1$ points from the unit simplex in \mathbb{R}^n , each corresponding to the diagonal of a matrix in \mathfrak{C}_{Het} , and sampling $3M_p/4 + 1$ points $\xi = (\xi_1, \dots, \xi_n)$, say, analogously, each

⁶³In the present context \mathfrak{M}_0^{lin} is spanned by the intercept. Thus, $I_1(\mathfrak{M}_0^{lin}) = \{1, \dots, n\}$ holds since $n \geq 2$. In general, to determine $I_1(\mathfrak{M}_0^{lin})$ numerically, the algorithm implemented in the R-package **hrt** (Preinerstorfer (2021)) first obtains a basis for \mathfrak{M}_0^{lin} , and then checks for every $i = 1, \dots, n$ whether or not the rank of the matrix obtained by appending the basis with $e_i(n)$ increases. This is done by a rank computation analogous to the one described in the last-but-one paragraph of Appendix E.3 of Pötscher and Preinerstorfer (2020), using the same function "rank" referred to there, and with tolerance parameter 10^{-8} .

point ξ giving rise to a diagonal of a matrix in \mathfrak{C}_{Het} via $(\xi_1^2, \dots, \xi_n^2) / \sum_{i=1}^n \xi_i^2$; (ii) trying all diagonal matrices with a single dominant coordinate 0.9999 and the other coordinates all equal to $0.0001/(n-1)$, so that the trace equals 1; (iii) $n^{-1}I_n$; (iv) using a maximizer of the quadratic form described in Remark E.3; and (v) using an additional initial value in case of a “large” critical value C as described in Remark E.4, making use of the conventions discussed in parentheses in that remark. This results in $M_0 = M_p + n + 2$ and possible one more (in case C is large) candidates for initial values.

2. M_1 was chosen as 500, the optimization algorithm run in Stage 1 was a constrained Nelder and Mead (1965) algorithm (the default in R’s “constrOptim” function), which was run with a relative tolerance parameter of 10^{-2} and a maximal number of iterations of $20n$.
3. M_2 was chosen as 1, the optimization algorithm run in Stage 1 was a constrained Nelder and Mead (1965) algorithm (the default in R’s “constrOptim” function), which was run with a relative tolerance parameter of 10^{-3} and a maximal number of iterations of $30n$.
4. DA (used by Algorithm 1) was run with the parameters “acc = 10^{-3} ” and “lim = 30000” using the function “davies” of the package **CompQuadForm**.

F.2 Details concerning Section 9.2

The smallest size-controlling critical values reported in Tables 4 and 5 in Section 9.2 were obtained by running Algorithm 3 (with algorithm A given by Algorithm 1 and a maximal number of 25 iterations in the while loop) as implemented in the R-package **hrt** (Preinerstorfer (2021)) version 1.0.0. Concerning A, the same input parameters as described in the enumeration at the end of Appendix F.1 were used but with $M_p = 500\,000$ (and with $n = 30$). Concerning Algorithm 3 we made the following choices for the required inputs:

1. $C_{low} = \max(C^*, C_{\text{hom}})$ is determined as follows: C_{hom} is determined by a line-search algorithm (using R’s uniroot function and monotonicity of the rejection probabilities in the critical value) with the rejection probabilities obtained from DA (in case $q = 1$) or via Monte Carlo, whereas C^* is determined as described in Appendix F.1. For more detail see Preinerstorfer (2021).
2. ϵ was set to 10^{-3} .

For computing the power functions in Section 9.2, we made use of (32) with the matrices A_C given in Lemma E.1 together with the implementation of the algorithm by Davies (1980) in the R-package **CompQuadForm** (Duchesne and de Micheaux (2010)) version 1.4.3 and with default parameters.

F.3 Additional figures for Section 9.2

The power functions for $n_1 = 9$ are given in Figure 4.

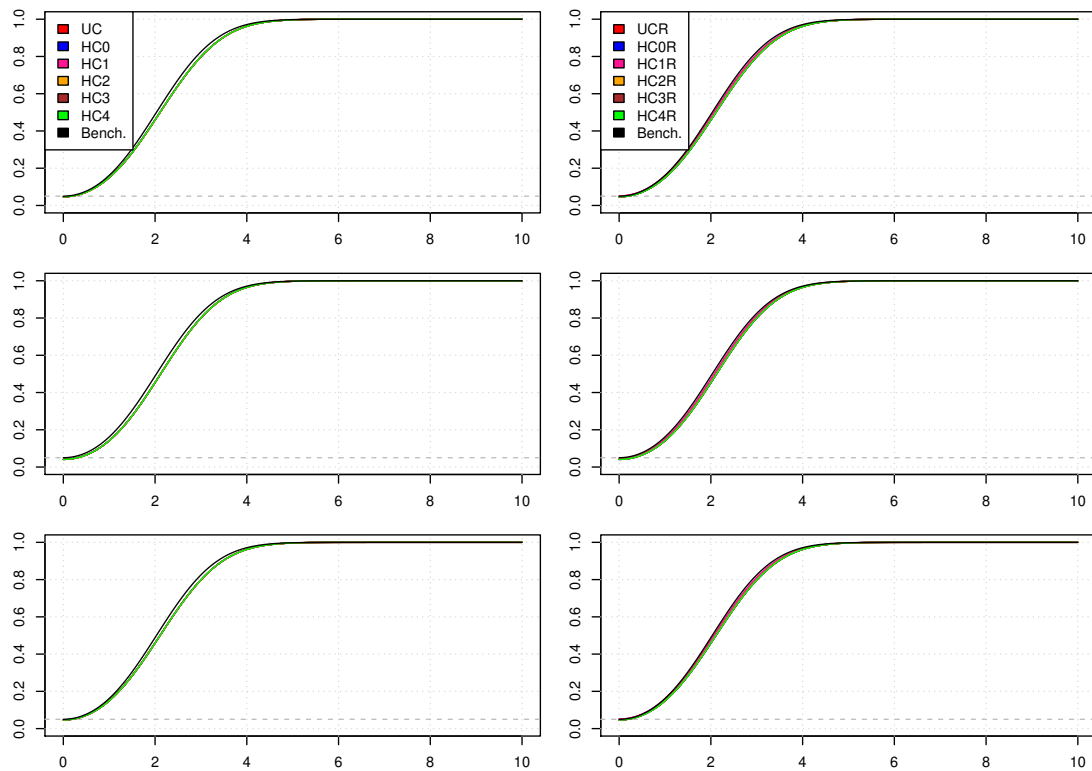


Figure 4: Power functions for $n_1 = 15$. Left column: tests based on unrestricted residuals (cf. legend). Right column: tests based on restricted residuals (cf. legend). The rows corresponds to Σ_a for $a = 1, 5, 9$ from top to bottom. The abscissa shows δ .

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