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Research Article

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Discriminating between (in)valid External Instruments and (in)valid Exclusion Restrictions

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Abstract: In models estimated by (generalized) method of moments a test on coefficient restrictions can either be based on a Wald statistic or on the difference between evaluated criterion functions. From their correspondence it easily follows that the statistic used for testing instrument validity, the Sargan-Hansen overidentifying restrictions (OR) statistic, is equivalent to an exclusion restrictions test statistic for a non-unique group of regressor variables. We prove that asymptotically this is the case too for incremental OR tests. However, we also demonstrate that, despite this equivalence of test statistics, one can nevertheless distinguish between either the (in)validity of some additional instruments or the (un)tenability of particular exclusion restrictions. This, however, requires to be explicit about the adopted maintained hypothesis. It also highlights that recent warnings in the literature that overidentifying restrictions tests may mislead practitioners should not be directed towards the test, but to practitioners who do not realize that inference based on such tests is unavoidably conditional on the validity of particular just-identifying statistically untestable assumptions.

Keywords: difference in Sargan-Hansen tests; exclusion restrictions; instrument validity tests; maintained hypothesis; overidentifying restrictions.

JEL Classifications: C01; C12; C26.

1 Introduction

In empirical econometric models it is often not unlikely that some explanatory variables are in fact endogenous, because their realizations are contemporaneously correlated with the error terms of the model. To overcome inference problems due to such correlation one often exploits external (non-explanatory) variables that should be uncorrelated with these errors. To verify whether these variables (instruments) seem uncorrelated with the disturbances indeed one can use a so-called test for overidentifying restrictions (OR), see Sargan (1958), Hansen (1982). The mandatory use of OR tests on a routine basis has been advocated at various places in the literature. See, for instance, the list mentioned in Baum et al. (2003, footnote 11). However, warnings that OR tests may mislead practitioners were recently issued in Parente and Santos Silva (2012) and Guggenberger (2012), and can also be found in, for instance, Hayashi (2000, p. 218) and Cameron and Trivedi (2005, p. 277).

In this study we will place these warnings regarding OR tests into perspective and highlight that they only apply if one is hesitant with respect to formulating an explicit maintained hypothesis, involving sufficient a priori orthogonality conditions and exclusion restrictions. Also we will expose the algebraic equivalences

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between OR and CR (coefficient restrictions) test statistics and explain why these tests nevertheless allow
distinct inferences. Moreover, we discuss to what degree similar issues arise for incremental OR tests.

For the sake of simplicity we will focus here mainly on the linear method of moments context, but our
derivations can be extended to more general settings, as we shall indicate. First, while introducing our nota-
tion, we review some general results on testing in a linear instrumental variables context, such as illustrat-
ing that the classical trinity of test principles established in the context of maximum likelihood inference,
constituted by Wald (W), likelihood ratio (LR) and Lagrange multiplier (LM), have their counterparts in a
semiparametric setting in which models can efficiently and robustly be estimated by (generalized) method
of moments, see also Newey and West (1987). In a linear method of moments context CR tests based on a
W statistic are in fact equivalent to the test obtained by taking the difference between evaluated criterion
functions (CF) for the restricted and the unrestricted model. The latter reminiscences of an LR-type test. By
employing an estimate for the disturbance variance obtained from the restricted model both W and CF can be
implemented as a kind of LM-type test. A very particular implementation of a CF test gives the Sargan-Hansen
statistic for testing overidentifying restrictions. Thus, the very same test statistic can be interpreted either as a
CR test or as an OR instrument validity test. However, we will highlight that these two interpretations require
different maintained hypotheses. This has not always been made very clear in earlier literature.

We also prove that in linear models similar equivalences and differences exist between exclusion restrictions
tests and so-called incremental overidentifying restrictions (IOR) tests (also addressed as difference in Sargan-
Hansen tests), which verify the validity of a subset of instruments. From these findings we conclude that strictly
following formal statistical test principles enables to characterize the context in which OR and IOR tests should
not mislead and in fact allow to distinguish inference regarding the (in)validity of some additional orthogonality
conditions from inference on the (un)tenability of particular exclusion restrictions. Practical problems do emerge
only when one is unable or unwilling to condition the analysis on a firm initial maintained hypothesis.

The structure of this study is as follows. In Section 2 we introduce the model and various test procedures
and concepts. Section 3 demonstrates how the standard Sargan test can either be used for testing instrument
validity or under a different maintained hypothesis can be interpreted as a CR test. Section 4 demonstrates to
what degree these results do apply to incremental Sargan tests as well. Finally Section 5 indicates options for
further generalizations and summarizes the conclusions.

2 Corresponding Test Principles

We focus on the single linear simultaneous regression model \( y = X \beta + \epsilon \) with \( \epsilon \sim (0, \sigma^2 I) \), where \( X = (x_{1}', \ldots, x_{j}')' \) is an \( n \times K \) matrix of rank \( K \) with unknown coefficient vector \( \beta \in \mathbb{R}^K \). To cope with \( E(x_i \epsilon) = \sigma_i \neq 0 \) for \( i = 1, \ldots, n \) we will employ an \( n \times L \) instrumental variables matrix \( Z = (z_1, \ldots, z_v)' \) of rank \( L \geq K \). The instrumental variables (IV)
or two-stage least-squares (2SLS) estimator is given by

\[
\hat{\beta} = (X'P_2X)^{-1}X'P_2y,
\]

where \( P_2 = A(A'A)^{-1}A' \) for any full column rank matrix \( A \). When \( E(x_i \epsilon) = 0 \) (the instruments are valid) and stand-
ard regularity conditions are fulfilled too (including sufficient relevance of the instruments) then the IV esti-
mator is consistent and asymptotically normal, whereas its variance can be estimated by \( \text{Var}(\hat{\beta}) = \hat{\sigma}^2(X'P_2X)^{-1} \),
where \( \hat{\sigma}^2 = \hat{\epsilon}'\hat{\epsilon}/n \) with residuals \( \hat{\epsilon} = y - X\hat{\beta} \).

A test on any set of \( K_j < K \) linear restrictions on the coefficients \( \beta \) can after a simple transformation of the
model be represented as testing exclusion restrictions \( H_{j'} \hat{\beta}_{j'} = 0 \) in

\[
y = X_1 \beta_1 + X_2 \beta_2 + \epsilon \sim (0, \sigma^2 I),
\]

where \( X_i \) is \( n \times K_i \) and \( \beta_j \) is \( K_j \times 1 \) for \( j = 1, 2 \). Making use of standard results on partitioned regression applied to
the second-stage regression, in which \( y \) is regressed on \( P_2X = (P_2X_1', P_2X_2') = (\tilde{X}_1', \tilde{X}_2') \), one easily finds that
\( \hat{\beta} = (\hat{\beta}_1', \hat{\beta}_2')' \), with
\[ \hat{\beta}_j = (\hat{X}_j' M_{\hat{X}_j} \hat{X}_j)^{-1} \hat{X}_j' M_{\hat{X}_j} y \]  

(3)

and \( \text{Var}(\hat{\beta}_j) = \sigma^2(\hat{X}_j' M_{\hat{X}_j} \hat{X}_j)^{-1} \), where \( M_x = I - P_{\hat{X}_j} \). The W statistic for testing \( \mathcal{H}_0: \beta_2 = 0 \) against alternative \( \mathcal{H}_1: \beta_2 \neq 0 \) readily follows and is given by

\[ W_{\beta_2} = \hat{\beta}_2 [\text{Var}(\hat{\beta}_2)]^{-1} \hat{\beta}_2. \]  

(4)

Under \( \mathcal{H}_0 \) it is asymptotically \( \chi^2(K_2) \) distributed, provided the maintained hypothesis does hold indeed. The maintained hypothesis (which is the union of \( \mathcal{H}_0 \) and \( \mathcal{H}_1 \) and all underlying assumptions) can here be characterized loosely as

\( \mathcal{M}_2: \)

(i) the actual data generating process (DGP) is nested in model (2) with \( K \) regressors;

(ii) \( X \) and \( Z \) show sufficient regularity;

(iii) \( n^{-1/2} Z' \varepsilon \overset{d}{\rightarrow} N(0, \sigma^2) \text{ plim } n^{-1} Z' Z \).

Substituting (3) in (4) and making use of a result on projection matrices for a partitioned full column rank matrix \( A = (A_1, A_2) \) which says

\[ \text{Var}(\hat{\beta}_2) = \text{Var}(\hat{\beta}_2) = \text{Var}(\hat{\beta}_2) = \text{Var}(\hat{\beta}_2), \]

we find

\[ \hat{\beta}_2 = (\hat{X}_j' M_{\hat{X}_j} \hat{X}_j)^{-1} \hat{X}_j' M_{\hat{X}_j} y \]

(6)

Hence, the test statistic can easily be obtained by taking the difference between the restricted and the unrestricted sum of squared residuals of second stage regressions, while scaling by a consistent estimator of \( \sigma^2 \). If the latter is obtained from the restricted residuals

\[ y' (M_{\hat{X}_j} - M_{\hat{X}_j}) y / \sigma^2, \]

(7)

which is asymptotically equivalent with \( W_{\beta_2} \).

A test statistic with similarities to the LR principle is found as follows. Estimator \( \hat{\beta}_2 \) is obtained by minimizing with respect to \( \beta \) a quadratic form in the vector \( Z'(y - X\hat{\beta}) \) in which a weighting matrix is used proportional to \( (Z'Z)^{-1} \). This yields a consistent estimator with optimal variance under the maintained hypothesis. Defining the criterion function as

\[ Q(\beta; y, X, Z, \sigma^2) = (y - X\beta)' Z(\hat{Z}' Z)^{-1} \hat{Z}' (y - X\beta) / \sigma^2, \]

(8)

one can verify that it attains its minimum for \( \hat{\beta} \) of (1), giving

\[ Q(\hat{\beta}; y, X, Z, \sigma^2) = (y - X\hat{\beta})' P_{\hat{Z}} (y - X\hat{\beta}) / \sigma^2 = (y' P_{\hat{Z}} y - 2\hat{\beta}' \hat{X} y + \hat{\beta}' \hat{X} \hat{X} \hat{\beta}) / \sigma^2, \]

which has in its numerator the difference between the sum of the squares of the second-stage residuals and the reduced form residuals. In the model under \( \beta_2 = 0 \), upon using the criterion function \( Q(\beta; y, X, Z, \sigma^2) \), one finds \( \hat{\beta}_1 \) of (6) with

\[ Q(\hat{\beta}_1; y, X, Z, \sigma^2) = (y' M_{\hat{X}_j} y - y' M_{\hat{X}_j} y) / \sigma^2, \]

The difference between these two minimized criterion functions is

\[ Q(\hat{\beta}; y, X, Z, \sigma^2) - Q(\hat{\beta}; y, X, Z, \sigma^2) = (y' M_{\hat{X}_j} y - y' M_{\hat{X}_j} y) / \sigma^2. \]

Now substituting either \( \hat{\sigma}^2 \) or \( \hat{\sigma}^2 \), we find that taking the difference between these two evaluated criterion functions yields either

\[ CF_{\hat{\beta}_j} = Q(\hat{\beta}; y, X, Z, \sigma^2) - Q(\hat{\beta}; y, X, Z, \sigma^2) = W_{\beta_2}. \]

(9)
or
\[ CF_{\beta_i}(\hat{\sigma}^2) = LM_{\beta_i}, \] (10)

These equivalences are algebraic, thus hold irrespective of the validity of \( M_i \).

### 3 The Sargan Test and Exclusion Restrictions

A special situation occurs in case \( L=K \). This specializes \( M_i \) to what we will indicate by \( M_i^{L=K} \). Under this maintained hypothesis the model is just identified, but overidentified under the \( K \) coefficient restrictions \( \beta = 0 \). When the model is just identified then, when minimizing \( Q(\beta; y, X, Z, \sigma^2) \), the \( L=K \) orthogonality conditions \( Z'(y-X\hat{\beta}) = 0 \) will all be satisfied in the sample, giving \( Q(\beta; y, X, Z, \sigma^2) = 0 \). So, in this particular situation, (9) specializes into

\[ W_{\beta_i}^{L=K} = CF_{\beta_i}^{L=K}(\hat{\sigma}^2) = Q(\beta_i; y, X_i, Z, \hat{\sigma}^2) = \hat{\varepsilon}'P\hat{\varepsilon} / \hat{\sigma}^2, \] (11)

and (10) into the asymptotically equivalent statistic

\[ LM_{\beta_i}^{L=K} = CF_{\beta_i}^{L=K}(\hat{\sigma}^2) = \hat{\varepsilon}'P\hat{\varepsilon} / \hat{\sigma}^2. \] (12)

The latter expression is the well-known Sargan test statistic which is used to check the validity of the instruments by testing the \( L-K_1 \) overidentifying restrictions in the model which has just the \( K_1 \) regressors \( X_\rho \), and can be expressed as

\[ y = X_1 \beta_1 + \varepsilon \] with \( \varepsilon \sim (0, \sigma^2 I). \) (13)

From its algebraic equivalence with a CR test it is immediately obvious that testing for OR has similarities with testing the validity of \( L-K_1 \) exclusion restrictions.

However, there are dissimilarities as well. Validity of the \( L \) orthogonality conditions \( E(z\varepsilon) = 0 \) is part of the maintained hypothesis \( M_i^{L=K} \) for the test of \( L-K_1 \) exclusion restrictions \( \beta_i = 0 \) in model (2), whereas the Sargan test is used to verify validity of the orthogonality conditions \( E(z\varepsilon) = 0 \) in model (13). Although using exactly the same test statistic and critical values, we will explicate that the two tests have intrinsically different \( \mathcal{H}_0 \) and \( \mathcal{H}_1 \) and build on different maintained hypotheses. We are not aware of literature where these differences and their consequences have been spelled out in full detail. Below we aim to fill this gap.

Note that expression (12) is invariant regarding \( X \). So, replacing \( X \) by any alternative \( n \times K_2 \) matrix \( X_2 \) would yield equivalent results for \( CF_{\beta_i}^{L=K}(\hat{\sigma}^2) \), provided \( P_z(X_1, X_2) \) has full column rank \( K=K_1 \). Therefore, \( (X_1, X_2) \) should have full column rank and the variables \( X_2 \) should be sufficiently related to the instruments \( Z \), but at the same time they may also be endogenous regarding \( \varepsilon \). Due to the projection of the regressors on the space spanned by the instruments \( Z \) in the second stage regression, and the presence of the regressors \( X_\rho \), the CR test actually does not test the explanatory power of \( X_2 \) as such, but just that of \( M_1 \), \( P_z(X_1, X_2) = (P_z - P_{X_1})X_2 \), which is its projection on the space spanned by \( Z \), as far as this is orthogonal to \( P_zX_1 \).

A viable choice for \( X_2 \) would be the following. Let \( F=F,(F,F) \) be a full rank \( L \times L \) matrix, where \( F \) is \( L \times L_1 \), \( F_2 \) is \( L_2 \times L_1 \), with \( L=L_1+L_2 \) and \( L=L_1+K_1 \). Now consider the transformed instruments \( Z'= (Z_1', Z_2') = ZF = (ZF_1, ZF_2) \), and let \( F \) be such that square matrix \( Z'X_1 \) has full rank. Obviously, \( F \) is nonunique. Note that if the transformed instruments \( Z_2' \) are not just relevant for \( X_1 \), but were valid for the errors \( \varepsilon \) of model (13) as well, then the coefficients \( \beta_1 \) would be just identified by them. When taking \( X_2 = M_{z'}Z_2' \), then \( P_z(X_1, X_2) \) has full column rank as required, and statistic \( CF_{\beta_1}^{L=K}(\hat{\sigma}^2) \) for testing \( \mathcal{H}_{0; \beta_1} = 0 \) in model

\[ y = X_1 \beta_1 + M_{z'}Z_2' \beta_2 + \varepsilon \] (14)

is therefore equivalent to the Sargan test statistic \( CF_{\beta_1}^{L=K}(\hat{\sigma}^2) \). Estimating \( \beta_1 \) of (14) by IV using the instruments \( Z \) or \( Z' \) (which is equivalent because \( P_z = P_{z'} \) ) yields
constitute valid \( \{0, \) plim \( \} \).

Thus, augmenting model (13) by \( K \) regressors \( M_{Z_1} Z_i' \) and using instruments \( Z \) or \( Z' \) yields coefficient estimates \( \hat{\beta}_1 \) for the regressors \( X \) which are equivalent to the simple IV estimator \( \hat{\beta}_1^* \) obtained in model (13) when just using the \( K \) instruments \( Z_i' \), irrespective of the validity of any of the instruments.

Apart from the interpretation of \( CF^{l-x}_K(\hat{\sigma}_1^2) \) as testing \( K_2=L-K \) exclusion restrictions \( \beta^\ast = 0 \) in model (2) under maintained hypothesis \( \mathcal{M}_2^{l-x} \), which presupposes the \( K \) orthogonality conditions \( E(\varepsilon|X) = 0 \), yet another coherent interpretation of the numerically equivalent test statistic \( CF^{l-x}_K(\hat{\sigma}_1^2) \) is the following. It tests whether the \( L-K \) variables \( Z_i' \) are additional valid instruments for model (13) under the maintained hypothesis given by \( \mathcal{M}_3^{l-x} \): (i) the DGP is nested in model (13) which has \( K_2<K=L \) regressors; (ii) \( X_i \) and \( Z \) (and thus \( Z' \)) show sufficient regularity; (iii) \( n^{-1/2} Z_i' \varepsilon \overset{d}{\rightarrow} N(0, \sigma^2_\varepsilon) \) plim \( n^{-1/2} Z_i' Z_i' \).

Note that for adopting \( \mathcal{M}_3^{l-x} \) in practice it might be desirable to specify matrix \( F_1' \); this issue will be addressed below.

Under \( \mathcal{M}_2^{l-x} \) the value of \( E(\varepsilon_i|\varepsilon_i') = \sigma_{\varepsilon_i} \) is unspecified. Obviously, if next to \( Z_i' \) the variables \( Z_i' \) are uncorrelated with \( \varepsilon \) then \( M_{Z_i} Z_i' \) should not have explanatory power in model (14). Hence, when adopting \( \mathcal{M}_2^{l-x} \), then under the null hypothesis \( \beta^\ast = 0 \) in model (14) all instruments \( Z \) are valid for model (13), implying \( \sigma_{\varepsilon_i} = 0 \) so that \( \hat{\beta}_1 \) is consistent, asymptotically normal and efficient for \( \beta^*_1 \). Below we will simply use the acronym AE (asymptotically efficient) to indicate for an estimator if, under the assumptions made, it is consistent with asymptotically normal distribution and smallest asymptotic variance (in a matrix sense).

The alternative \( \beta^* \neq 0 \) implies \( \sigma_{\varepsilon_i} \neq 0 \). So, in the context of model (14) test statistic \( CF^{l-x}_K(\hat{\sigma}_1^2) \) tests the null hypothesis given by the \( L-K \) orthogonality conditions \( E(\varepsilon_i|X) = 0 \) or overidentifying restrictions \( \beta^\ast = 0 \). When these conditions/restrictions are invalid then, given validity of \( \mathcal{M}_2^{l-x} \), estimator \( \hat{\beta}_1 = (Z_i' X_i)^{-1} Z_i' y \) for \( \beta^\ast_1 \) is AE in model (13). In the extended regression model (14) all variables in both \( Z_i' \) and \( Z_i' \) constitute valid instruments regarding \( \varepsilon \) under \( \mathcal{M}_3^{l-x} \), because inclusion of the regressors \( M_{Z_i} Z_i' \) purges the disturbances from their correlation with \( Z_i' \). That this regression too produces the AE estimator \( \hat{\beta}_1 \) for the coefficients \( \beta^\ast_1 \), as is proved by (15), is because it constitutes a special case of the so-called “partialling out” result which will be reviewed at the end of the next section.

It is the case that truth of the null \( \beta^\ast_2 = 0 \) under \( \mathcal{M}_2^{l-x} \) implies \( E(\varepsilon_i|X) = 0 \), and so does truth of the null \( E(\varepsilon_i|X) = 0 \) under \( \mathcal{M}_3^{l-x} \). Under both null hypotheses estimator \( \hat{\beta}_1 \) is AE for \( \beta^\ast_1 \). However, imposing the respective null hypotheses implies either accepting coefficient restrictions or accepting extra orthogonality assumptions. Invalidity of \( \beta^\ast_2 = 0 \) under \( \mathcal{M}_2^{l-x} \) renders \( \hat{\beta}_1 \) the AE estimator for \( \beta^\ast_1 \), but invalidity of \( E(\varepsilon_i|X) = 0 \) under \( \mathcal{M}_3^{l-x} \) renders \( \hat{\beta}_1 \) AE for the coefficients of just the regressors \( X_i \) since in this setup the DGP is supposed to be nested in model (13) which has just \( K \) regressors.

So, on the basis of the single test statistic \( LM^{l-x}_K(\hat{\sigma}^2) = CF^{l-x}_K(\hat{\sigma}^2) \), or the asymptotically equivalent \( W^{l-x}_K(\hat{\sigma}^2) \), one can discriminate between either (in)validity of \( L-K \) exclusion restrictions in model (2) or (in)validity of \( L-K \) external instruments in model (13). This just hinges upon which of the nonnested maintained hypotheses has actually been adopted: \( \mathcal{M}_2^{l-x} \) or \( \mathcal{M}_3^{l-x} \). In both contexts the tests concern a possible loss/gain in the degree of overidentification by \( L-K \). Namely, either due to rejecting/accepting exclusion restrictions on the coefficients of potential explanatory variables \( X_i \) in model (2), or on accepting/rejecting exclusion restrictions on the coefficients of the auxiliary regressors \( M_{Z_i} Z_i' \) in model (14). The latter implies accepting/rejecting validity of \( L-K \) external instruments \( Z_i' \).

The warnings of Parente and Santos Silva (2012) and Guggenberger (2012) concern the problems that emerge when using the Sargan statistic for testing jointly the \( L \) orthogonality conditions \( E(\varepsilon_i|X) = 0 \) in model
for model (13) can be tested by a so-called incremental Sargan statistic given in (1) and ensuring just identification of $\beta^*_i$.

It is the researcher who may decide whether (s)he leaves these $K_i$ orthogonality conditions implicit, or makes them explicit by actually specifying the matrix $F_i$, for instance by choosing $F_i=(I, O)'$ and then thus simply assuming that the first $K_i$ instruments in the matrix $Z$ are valid in model (13). In that special case the explicit null hypothesis of the Sargan test is that the instruments $M_{Z_i}Z_i$ are valid too, which would imply $E(z_i'\epsilon_i)=0$ in model (13).

### 4 The Incremental Sargan Test and Exclusion Restrictions

The validity of a subset of the instruments $Z_i'$ for model (13) can be tested by a so-called incremental Sargan or IOR test. For that we will examine too in what way such a procedure has (dis)similarities with testing exclusion restrictions. However, for this analysis we will take as our starting point the usual model with $K$ regressors and $L>K$ instruments and therefore have to introduce some further notation.

We consider a partition of the matrix of $L$ instruments, denoted as $Z=(Z_m, Z_a)$, where for $j\in\{m, a\}$ matrix $Z_j=(z_j, \ldots , z_{jm})'$ is $n\times L_j$. Matrix $Z_m$, with $L_m\geq K$, contains the instruments which are maintained to be valid, whereas the validity of the additional $L_a>0$ instruments $Z_a$ will be examined by testing. So the maintained hypothesis is now given by

$$\mathcal{M}_{L_a}^{1-K}: (i) \text{ the DGP is nested in model } y=X\beta^*+\epsilon, \text{ which has } K \text{ regressors}; (ii) X \text{ and } Z=(Z_m, Z_a) \text{ with } L>L_m\geq K \text{ show sufficient regularity}; (iii) n^{-1/2}Z_m'\epsilon \sim N(0, \sigma^2 I) \text{ plim } n^{-1/2}Z_a'Z_m.$$

Just using the instruments $Z_m$ yields estimator and residuals

$$\hat{\beta}_m=(X'P_{z_m}X)^{-1}X'P_{z_m}y, \quad \hat{\epsilon}_m=y-X\hat{\beta}_m.$$  

From the foregoing section it follows that the Sargan statistic

$$S_m=n\hat{\epsilon}_m'P_{z_m}\hat{\epsilon}_m/\hat{\epsilon}_m'\hat{\epsilon}_m$$

is asymptotically $\chi^2(L_m-K)$ distributed under $\mathcal{M}_1^{1-K}$ when $L_m> K$, whereas it is zero for $L_m=K$. Using the additional instruments as well yields $\hat{\beta}$ given in (1) and $\hat{\epsilon}=y-X\hat{\beta}$ and enables to calculate

$$S_m=n\hat{\epsilon}'P_{z_m}\hat{\epsilon}/\hat{\epsilon}'\hat{\epsilon}.$$  

Now the null hypothesis $E(z_i'\epsilon_i)=0$, which involves $L_a$ orthogonality conditions for model (2) additional to the $L_m$ orthogonality conditions already implied by $\mathcal{M}_1^{1-K}$, can be tested by the incremental Sargan statistic

$$IS_a=S_m-S_a,$$

which under $\mathcal{M}_1^{1-K}$ has asymptotic null distribution $\chi^2(L_a)$. Note that for $L_m=K$ statistic $IS_a$ corresponds to the Sargan test examined in the foregoing section, but now the role of $X_1$ is played by $X$, so $K_1$ is now $K$, whereas $K_2$ is given here by $L$, and what earlier was called $\hat{\epsilon}$ is given by $\hat{\epsilon}$ now.

Most published derivations of the null distribution of $IS_a$ start off from a generalized method of moments (GMM) context. Our simple linear framework allows a very concise derivation, which goes as follows. Without loss of generality we assume $Z_m'Z_a=0$, which may have been achieved by replacing the original $Z_a$ by $M_{Z_m}Z_a$. $P_z$ is not affected by this partial orthogonalization, but now $P_z=P_{z_m}+P_z$. 


From \( P \hat{e} = P_1 [ I - X (X_1' P_1 X_1 X_1')^{-1} X_1' ] e = (P_1 - P_\epsilon) e \) we find
\[
\hat{e}' P \hat{e} = e' P_1 e - e' P_\epsilon e = (P_1 + P_\epsilon - P_1 P_\epsilon P_1) e.
\]
The matrix in this quadratic form is symmetric and idempotent. The latter follows from \( P_1 P_\epsilon P_1 = 0 \) and \( (P_1 + P_\epsilon P_1) P_\epsilon P_1 = (P_1 + P_\epsilon) X (X_1 P_1 X_1')^{-1} X_1 P_\epsilon P_1 = P_1 P_\epsilon P_1 \).

Since \( tr(P_1 + P_\epsilon P_1) = 0 \) and \( (P_1 + P_\epsilon P_1) P_\epsilon P_1 = (P_1 + P_\epsilon) X (X_1 P_1 X_1')^{-1} X_1 P_\epsilon P_1 = P_1 P_\epsilon P_1 \), this means that under \( \mathcal{M}_{L-K}^1 \) and the null hypothesis \( E(z_i e) = 0 \) \((i = 1, \ldots, n)\), so when all instrument in \( Z \) are valid, \( IS_{a} \rightarrow \hat{\epsilon}' \hat{\epsilon} / n \sim \sigma^2 \).

We shall now examine tests for omitted regressors (or exclusion restrictions) and seek cases which show correspondences with statistic \( IS_{a} \). Consider the extended model with \( K_a \leq L-K \) additional regressors \( X_{a} \) given by

\[
y = X \beta + X_a \beta_a + \epsilon_a, \quad \text{with} \quad \epsilon_a \sim (0, \sigma^2 I),
\]
and the maintained hypothesis defined by

\( \mathcal{M}_{L-K}^1 : (i) \) the DGP is nested in model \( y = X \beta + X_a \beta_a + \epsilon_a \) which has \( K + K_a \leq L \) regressors; \( (ii) \) \( X, X_a \) and \( Z \) show sufficient regularity; \( (iii) \) \( n^{-1/2} Z \epsilon \rightarrow N(0, \sigma^2) \) plim \( n^{-1/2} Z' Z \).

Let us indicate the IV results when estimating model (20) using instruments \( Z \) as \( y = X \hat{\beta} + X_a \hat{\beta}_a + \hat{\epsilon}_a \). Then the LM-like test statistic for \( H_0: \beta_a = 0 \) is given by

\[
CF_{\beta_a} (\hat{\sigma}_a^2) = (\hat{\epsilon}' P_\epsilon \hat{\epsilon} - \hat{\epsilon}_a' P_\epsilon \hat{\epsilon}_a) / \hat{\sigma}_a^2,
\]
which under \( \mathcal{M}_{L-K}^1 \) has asymptotic null distribution \( \chi^2(K_a) \).

Still assuming without loss of generality that \( Z_a' Z_a = O \), we shall first examine the specific case \( X_a = Z_a \) for which expression (21) can be specialized by using the following results. Substituting \( X_a = Z_a \) we obtain

\[
\hat{\beta}_a = (X P_1 M_{a_1} P_a X_a)^{-1} X P_1 M_{a_1} y = (X P_1 M_{a_1} y |
\]
because \( M_{a_1} P_a = P_a - P_{a_1} = P_a - P_{a_1} = P_a \). And, since in IV estimation the residuals are orthogonal to the second-stage regressors, we have \( Z_a' \epsilon_a = 0 \). This yields \( P_\epsilon \hat{\epsilon}_a = (P_a + P_{a_1}) \hat{\epsilon} = P_a (y - X \hat{\beta}_a - Z \hat{\beta}_a) = P_a (y - X \hat{\beta}_a) = P_a (y - X \hat{\beta}_a) = P_a \hat{\epsilon}_a \), where we used (22) for the last equality. From this we find

\[
CF_{\beta_a} (\hat{\sigma}_a^2) = (\hat{\epsilon}' P_\epsilon \hat{\epsilon} - \hat{\epsilon}_a' P_\epsilon \hat{\epsilon}_a) / \hat{\sigma}_a^2 = (\hat{\epsilon}' P_\epsilon \hat{\epsilon} - \hat{\epsilon}_a' P_{a_1} \hat{\epsilon}_a) / \hat{\sigma}_a^2 = S_{ma} - S_{ma}, \quad \hat{\sigma}_a^2 \]

Under \( \mathcal{M}_{L-K} \) and \( H_0: \beta_a = 0 \) both \( \hat{\epsilon}_a' \hat{\epsilon}_a / n \) and \( \hat{\epsilon}' \hat{\epsilon} / n \) converge to \( \sigma^2 \), thus the ratio \( \hat{\epsilon}_a' \hat{\epsilon}_a / \hat{\epsilon}' \hat{\epsilon} \) equals 1 asymptotically. Therefore, under their respective null hypotheses, the incremental Sargan test statistic \( IS_{a} = S_{ma} - S_{ma} \) is asymptotically equivalent to the exclusion restrictions test statistic \( CF_{\beta_a} (\hat{\sigma}_a^2) \).

In fact, this also holds for more general matrices than just \( X_a = Z_a \). Consider instead

\[
X_a = Z_a C_a + X C_a + M_a C_a + \epsilon C_a,
\]
where the finite matrices \( C_a, C_a, C_a, C_a \) are \( L_a \times K_a, K_a \times K_a, n \times K_a \) and \( 1 \times K_a \), respectively. Now writing \( \hat{X}_a = P_a X_a \) and \( \hat{X} = P_a X \) it follows that \( \hat{X}_a = Z_a C_a + \hat{X} C_a + P_a \epsilon C_a \). Under the null hypothesis \( \beta_a = 0 \), we have plim \( n^{-1/2} Z \epsilon = 0 \), hence \( P_a \epsilon C_a = 0 \), thus \( M_a \hat{X}_a = (M_a X_a C_a \epsilon C_a) \). The regularity assumption of \( \mathcal{M}_{L-K} \) requires this to have full column rank, hence \( C_a \) should have rank \( K_a \leq L_a \). If \( K_a = L_a \) and \( C_a \) has full rank we obtain \( P_{M_a} \hat{X}_a = P_{M_a} \hat{X} \). Thus, tests for the omission of regressors \( X_a \) which belong to group (24) with \( C_a \) of full rank \( K_a = L_a \) will yield test statistics which under the null are all asymptotically equivalent with statistic \( IS_{a} \).
Estimator $\hat{\beta}$ is AE both under $\mathcal{M}^{(1)}$ with valid exclusion restrictions $\beta_0 = 0$ and under $\mathcal{M}^{(2)}$ with valid additional instruments $Z_a$. However, under $\mathcal{M}^{(1)}$ while instruments $Z_a$ are invalid $\hat{\beta}_m$ is AE, whereas under $\mathcal{M}^{(2)}$ with $\beta_0 \neq 0$ the estimators $\hat{\beta}_1$ and $\hat{\beta}_m$ are AE in model (20), yielding an estimator for $\beta$ asymptotically equivalent to $\hat{\beta}_m$ only if $X_0$ belongs to group (24) with $K = L_1, C$ of full rank and $C = 0$.

The algebraic estimator equivalence for $X_a = Z_a$ given in (22) reestablishes the so-called partialling out result, which says that IV coefficient estimates are invariant when (putative) exogenous regressors are deliberately omitted, provided they are also removed from and filtered out of the remaining instruments. In Hendry (2011) this analytic result is established by simulation, which leads to the unsatisfactory conclusion that (in our notation) $\hat{\beta}_1$ and $\hat{\beta}_m$ “hardly differ.” Moreover, in the simulation design used the instruments are all valid, $L = 3, K = 2, L_a = 1$ and all variables are normally distributed, whereas neither of these is required for the algebraic result to hold exactly.

5 Interpretation and Conclusion

The above results lead to clear guidelines only when a researcher is willing to adopt unambiguously a particular maintained hypothesis, which (s)he should according to sound Neyman-Pearson statistical test methodology. In practice, however, a more flexible though opportunistic approach may often seem more appealing. The two juxtaposed maintained hypotheses $\mathcal{M}^{(1)}$ and $\mathcal{M}^{(2)}$ in Section 3, and their generalizations $\mathcal{M}^{(1)}$ and $\mathcal{M}^{(2)}$ in Section 4, both assume that knowledge is available already of a model specification in which the DGP is nested, whereas endeavors to reach that stage form usually the major challenge in an empirical modeling exercise. This explains why researchers when obtaining a significant OR test may either conclude: (a) some instruments are invalid, (b) the model specification is deficient (so its implied errors may correlate even with instruments which would be valid for an adequate model specification), or (c) both model and instruments are unsound. We have shown that this shilly-shally is a consequence of not building on a firm maintained hypothesis. Under $\mathcal{M}^{(2)}$ an insignificant OR statistic is either due to validity of the tested additional moment conditions, or to lack of power of the test under (a), but not to (b) or (c), whereas a significant OR statistic is either due to a Type I error or to invalid instruments. Practitioners who interpret an insignificant OR statistic as approval of the chosen model specification and the employed instruments should always realize that this presupposes validity of the untested hypothesis that this instrument set contains a subset which already just-identifies the coefficients of this chosen model specification.

The presented results on instrument validity and coefficient restriction tests can rather straightforwardly be generalized to linear models with nonspherical disturbances estimated by generalized method of moments, because GMM conforms to IV/2SLS after proper transformations. When $\Sigma = (0, \sigma^2 \Omega)$ this transformation requires (as in GLS) to premultiply the model by an $\times n$ matrix $\Psi'$, where $\Psi' \Psi = \Omega^{-1}$, so that $\epsilon^* = \Psi \epsilon - (0, \sigma^2 I)$, but employ now as instruments $Z^* = (\Psi')^{-1} Z$, see Kiviet and Feng (2016). Hence, all above results can easily be reformulated for the linear GMM context, where (incremental) OR tests are usually addressed as Sargan-Hansen tests. In case of unknown heteroskedasticity (hence $\Omega$ is diagonal but not available), rather than robustifying inefficient IV, its findings $\hat{\beta}$ and $\hat{\epsilon} = y - X \hat{\beta}$ can be employed in a second step to obtain asymptotically efficient feasible GMM inference. This can be represented as follows. Let $y_i = y_i / \hat{\epsilon}_i$, $X^* = (x_{1i}, \ldots, x_{ni})'$ with $x_{1j} = x_{ij} / \hat{\epsilon}_i$ and $Z^* = (z_{1i}, \ldots, z_{ni})'$ with $z_{1j} = z_{ij} / \hat{\epsilon}_i$. Next regressing $y'$ on $X'$ using instruments $Z'$ yields $\hat{\beta}' = (X' P_{2} X')^{-1} X' P_{2} y' = [X' Z(Z' Z)^{-1} Z' X]^{-1} X' Z(Z' Z)^{-1} Z' y'$, with $Z' Z = \Sigma_{ii} = z_{ij} z_{ij}$. Its variance could be estimated consistently by $\hat{\text{Var}}(\hat{\beta}') = [X' Z(Z' Z)^{-1} Z' X]^{-1}$, and writing $\hat{\epsilon}' = y' - X' \hat{\beta}'$ the

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1 See, for instance, Hayashi (2000, p. 218) and Cameron and Trivedi (2005, p. 277).
2 When embedded into a properly extended framework generalizations for nonlinear models can be obtained as well; see, for instance, Newey (1985).
3 See Newey and West (1987, p. 786).
Sargan-Hansen statistic is $S' = \bar{z}' P_x \bar{z} = (y - X \hat{\beta}')(Z'Z)^{-1}Z'(y - X \hat{\beta})$, which is asymptotically equivalent to the GMM maximized criterion function. Hence, also for feasible GMM all earlier findings on CR and IOR tests have their (asymptotically equivalent) counterparts. Note, however, that a robustified variance estimator for $\hat{\beta}$ would be $\bar{z}'(\sum_{i=1}^n \bar{z}_i \bar{z}_i')^{-1}\bar{z} = (y - X \hat{\beta}')(Z'Z)^{-1}Z'(y - X \hat{\beta})$. Due to the consistency of $\hat{\beta}$ this is asymptotically equivalent to the two-step statistic $S'$, and thus a robustified CF test has correspondences with a CR test based on a two-step feasible GMM estimator $\hat{\beta}'$ and no longer with one based on the inefficient $\hat{\beta}$.

Also in the extended contexts indicated in the preceding paragraph the practical relevance of the here presented findings is the following. An (incremental) Sargan-Hansen test on the validity of (particular) orthogonality conditions uses exactly the same (or at least an under the null asymptotically equivalent) test statistic with the same asymptotic null distribution as a test for zero restrictions on the coefficients of additional regressor variables which may stem from a particular fairly wide group. However, in essence these two test procedures build on different nonnested maintained hypotheses which enables them to produce discriminative inferences.

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**References**


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4 Baum et al. (2003, p. 18) claim that $S'$ and $S'$ are even numerically equivalent, which does not seem right.