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**THE OPTIMAL SIZE OF A PRELIMINARY TEST  
OF LINEAR RESTRICTIONS IN A  
MIS-SPECIFIED REGRESSION MODEL**

**David A. E. Giles, Offer Lieberman & Judith A. Giles**

***Discussion Paper***

**No. 9006**

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David E.A. Giles

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Abstract

When the choice of estimator for the coefficients in a linear regression model is determined by the outcome of a prior test of the validity of restrictions on the model, Brook (1976) has shown that a mini-max (risk) regret criterion leads to the simple rule that the optimal critical value for the preliminary test is approximately two in value, regardless of the degrees of freedom. We show that this result no longer holds in the (likely) event that relevant regressors are excluded from the model at the outset.

Key Words : Preliminary Test; Conditional Inference; F-Test; Optimal Size.

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## 1. INTRODUCTION

Applied statisticians and econometricians routinely adopt a "preliminary - test" estimator (PTE) when using the linear multiple regression model. That is, they search for the preferred specification of the model by testing exact (and often "zero") linear restrictions on the coefficients, and then apply Ordinary Least Squares (OLS) or Restricted Least Squares (RLS) estimation according to the outcome of the test. (For example, see Bancroft (1944), Brook (1976), Judge & Bock (1978)). Of course, other pre-test regression strategies are also common, such as those associated with testing for homogeneity of the error variance (e.g. Toyoda and Wallace (1975), Bancroft and Han (1983)) and testing for serial independence of the errors (e.g., King and Giles (1984)).

The finite sample risks (say, under quadratic loss) of PTEs are complicated functions of all aspects of the problem, including the chosen size (and hence critical value(s)) for the two "component" estimators, and are discontinuous functions of the sample data. Accordingly, they are generally inadmissible (e.g., Cohen (1965)). However, PTEs are routinely used, often without an appreciation that their sampling properties differ crucially from those of the component estimators, and so they warrant close scrutiny.

Given that the risk of a PTE depends, in part, on the pre-test size, various criteria have been suggested for choosing the "optimal" size. In the case of the PTE arising from the testing of exact linear restrictions on the coefficients of the linear model, two such optimality criteria are those suggested by Brook (1976) and Toyoda and Wallace (1976).

A recent development in the literature on regression PTEs has been an investigation of the extent to which their properties are affected if the model is mis-specified in various ways (e.g., Ohtani (1983), Mittelhammer (1984), Giles (1986), Giles and Clarke (1989), Giles (1990)). To date, there has been no discussion of the effects that model mis-specification may have on

the "optimal" choice of pre-test size. This paper addresses this question by extending Brook's analysis to the important case where relevant regressors are omitted from the model.

## 2. THE MODEL AND DEFINITIONS

Suppose that the data-generating process is

$$y = X\beta + Z\gamma + u; \quad u \sim N(0, \sigma^2 I) \quad (2.1)$$

where  $X$  and  $Z$  are  $(n \times k)$  and  $(n \times g)$  respectively, each of full rank and non-stochastic. However, the model fitted to the data is

$$y = X\beta + \varepsilon, \quad (2.2)$$

where  $\varepsilon = Z\gamma + u = \xi + u$ , say. Within the framework of the (mis-specified) model (2.2), we test  $m$  independent linear restrictions:

$$H_0 : R\beta = r \quad \text{vs} \quad H_A : R\beta \neq r,$$

where  $R$  is  $(m \times k)$ , of rank  $m$ ;  $r$  is  $(m \times 1)$ ; and both are non-stochastic. Being unaware that the model is mis-specified, we use the usual "F-statistic",

$$f = (r - R\tilde{\beta})' (RS^{-1}R')^{-1} (r - R\tilde{\beta}) / ms^2,$$

where  $S = X'X$ ;  $\tilde{\beta} = S^{-1}X'y$  is the OLS estimator of  $\beta$ ;  $s^2 = y'(I - XS^{-1}X')y/\nu$ ; and  $\nu = (n - k)$ .

It is readily shown (e.g. Mittelhammer (1984)) that  $f$  is  $F_{(m, \nu; \lambda_n, \lambda_d)}''$ ; that is, doubly non-central F with  $m$  and  $\nu$  degrees of freedom, and non-centrality parameters

$$\lambda_n = (RS^{-1}X'\xi - \delta)'(RS^{-1}R')^{-1}(RS^{-1}X'\xi - \delta)/2\sigma^2$$

$$\lambda_d = \xi'(I - XS^{-1}X')\xi/2\sigma^2,$$

where  $\delta = R\beta - r$ .

The PTE of  $\beta$  is based on  $\tilde{\beta}$ ; the RLS estimator of  $\beta$ ,  $\beta^* = \tilde{\beta} - S^{-1}R'(RS^{-1}R')^{-1}(R\tilde{\beta} - r)$ ; and the use of  $f$  to test  $H_0$ :

$$\hat{\beta} = \begin{cases} \tilde{\beta} & \text{if } f > c(\alpha) \\ \beta^* & \text{if } f \leq c(\alpha) \end{cases}$$

where  $c(\alpha)$  is the size- $\alpha$  critical value for the central F statistic with  $m$  and  $\nu$  degrees of freedom.

The sampling properties of  $\tilde{\beta}$ ,  $\beta^*$  and  $\hat{\beta}$  will be compared on the basis of (relative) predictive risk under quadratic loss. For any estimator,  $b$ , of  $\beta$ , this is defined as  $\rho(Xb, E(y)) = E\left[\left(Xb - E(y)\right)' \left(Xb - E(y)\right)\right]/\sigma^2$ , which is equivalent to the risk of  $b$  itself with orthonormal regressors. So,

$$\rho(X\tilde{\beta}, E(y)) = (k + 2\lambda_d) \quad (2.3)$$

$$\rho(X\beta^*, E(y)) = (k + 2\lambda_d + 2\lambda_n - m) \quad (2.4)$$

$$\rho(X\hat{\beta}, E(y)) = (k + 2\lambda_d + (4\lambda_n - m)P_2 - 2\lambda_n P_4) \quad (2.5)$$

where  $P_i = \Pr\left[F_{(m+i, \nu; \lambda_n, \lambda_d)}'' \leq \frac{cm}{m+i}\right]$ . See Mittelhammer (1984) and Giles (1990).

If  $\xi = 0$ , then  $\lambda_d = 0$ , the fitted model is correctly specified, and (2.3) - (2.5) collapse to the corresponding expressions given by Brook (1976). Note that  $\rho(X\tilde{\beta}, E(y)) = \rho(X\beta^*, E(y))$  when  $\lambda_n = m/2$ , regardless of the value of  $\lambda_d$ . Also, from (2.3)-(2.5), it is readily shown for any  $\lambda_d$  (i.e. degree of model mis-specification), that  $\rho(X\hat{\beta}, E(y)) = \rho(X\tilde{\beta}, E(y))$  for some  $\lambda_n \in (m/4, m/2)$ ; and



that  $\rho(\hat{X}\hat{\beta}, E(y))$  has a unique mode at a value of  $\lambda_n$  greater than that for which  $\rho(\hat{X}\hat{\beta}, E(y)) = \rho(X\beta^*, E(y))$ . This is illustrated in Figure 1 for a specific degree of mis-specification. As Mittelhammer (1984) notes, the OLS, RLS and PTE risks are unbounded as  $\lambda_d \rightarrow \infty$  (for a given  $\lambda_n$ ). Further, for fixed  $c$  and  $\lambda_n$ ,  $\rho(\hat{X}\hat{\beta}, E(y)) \rightarrow \rho(X\beta^*, E(y))$  as  $\lambda_d \rightarrow \infty$ .

### 3. OPTIMAL SIZE OF THE TEST

Clearly, from (2.5), the predictive pre-test risk depends on  $c = c(\alpha)$ . For any particular value of  $\lambda_d$  the predictive risks, as functions of  $\lambda_n$ , have the same essential characteristics as when  $\lambda_d = 0$ . In particular, of the three estimators considered, RLS is preferred if  $\lambda_n < m/2$ , and OLS is preferred if  $\lambda_n > m/2$ . We can define the "regret" associated with using the PTE as

$$R(\hat{X}\hat{\beta}) = \begin{cases} \rho(X\hat{\beta}) - \rho(X\beta^*) & ; \quad \lambda_n < m/2 \\ \rho(X\hat{\beta}) - \rho(X\tilde{\beta}) & ; \quad \lambda_n \geq m/2. \end{cases}$$

As is well known, if  $\lambda_d = 0$ , then for any  $c$  there is a unique  $R^L = \sup_{\lambda_n < m/2} R(\hat{X}\hat{\beta})$  and a unique  $R^U = \sup_{\lambda_n \geq m/2} R(\hat{X}\hat{\beta})$ . Further, an increase in  $c$  leads to a decrease in  $R^L$  and an increase in  $R^U$ . The same result applies for any fixed  $\lambda_d > 0$ . So for a given  $\lambda_d$ , an optimal choice of  $c$  may be defined as  $c = c^*$  such that  $R^L = R^U$ , and both regrets are simultaneously minimized. This is the "mini-max - regret" rule adopted by Brook (1976) (and similar to that of Sawa and Hiromatsu (1973)) for the properly specified model ( $\lambda_d = 0$ ). The computations needed to obtain  $c^*$ , for any  $\lambda_d$ , are equivalent to those needed in Brook's case, but with doubly non-central F probabilities replacing his non-central F probabilities.

#### 4. RESULTS

Optimal critical values,  $c^*$ , are reported in Table 1 for several values of  $\lambda_d$ . These were calculated using a FORTRAN program written by the authors and executed on a VAX8350. The program incorporates Davies' (1980) algorithm to evaluate the doubly non-central F probabilities. The chosen degrees of freedom match those considered by Brook, and his results correspond to  $\lambda_d = 0$ . (There are some minor differences between our results and his for small and large degrees of freedom. These can be accounted for if one adopts less stringent convergence tolerances than we have used in our program.) The optimal significance levels ( $\alpha^*$ ) corresponding to  $c^*$ , and based on the central  $F_{m,\nu}$  distribution, are reported in Table 1. These are also computed using Davies' algorithm.

The strongest feature of Brook's results for the properly specified model is that  $c^*$  is always close to two in value, regardless of the degrees of freedom. (Using a different criterion, Toyoda and Wallace (1976) concur with this result for  $m \geq 5$ .) As Figure 1 and Table 1 show, this result is undermined if the model is mis-specified through the omission of regressors. In this case  $c^*$  is sensitive to the degrees of freedom and can differ substantially from the values suggested by Brook. In addition, for any  $m$  and  $\nu$ ,  $c^*$  declines monotonically as the degree of model mis-specification increases. Accordingly, the optimal pre-test size ( $\alpha^*$ ) increases monotonically with  $\lambda_d$ , for fixed degrees of freedom. This accentuates the other strong feature of the results obtained by Brook (and Toyoda and Wallace (1976)) - the optimal pre-test size is frequently much greater than the commonly assigned sizes of 1% or 5%.

## 5. CONCLUDING DISCUSSION

We have focussed attention on pre-testing in the context of a model which is under-specified. The case where extraneous regressors are included in the model does not require separate consideration. Giles (1986) shows that all of the usual risk results hold (as in Figure 1, with  $\lambda_d = 0$ ) in this case with a simple re-definition of  $\lambda_n$ . It follows that Brook's results (that is, the results in our Table 1 for  $\lambda_d = 0$ ) apply directly to the case of over-fitted models.

The latter results, which suggest that  $c^* = 2$  (approximately) regardless of the degrees of freedom, have obvious practical appeal. They offer a simple prescription to be followed in empirical work. However, as our results show, this prescription is dangerously misleading if the model is under-specified. It would be helpful to have a substitute prescription in the face of possible such model mis-specification. Table 1 does not provide this, given that  $\lambda_d$  is generally unknown.

If an upper bound,  $\bar{\lambda}_d$ , could be placed on  $\lambda_d$ , then the following generalised mini-max regret criterion might be considered: for some value of  $c$ , take a sequence of  $R^L$  values for  $\lambda_d \in [0, \bar{\lambda}_d]$ , and a corresponding sequence of  $R^U$  values. Then, vary  $c$  to  $c^{**}$ , say, to equate  $\sup_{0 < \lambda_d < \bar{\lambda}_d} (R^U)$  and  $\sup_{0 < \lambda_d < \bar{\lambda}_d} (R^L)$ . It is readily shown that  $c^{**}$  is unique. The implications of such a criterion are illustrated in Table 2, where  $\alpha^{**}$  is the test size corresponding to  $c^{**}$ , based on the central F distribution. The difficulty is that if  $\bar{\lambda}_d$  is unknown, then  $c^{**} \rightarrow 0$  as  $\bar{\lambda}_d \rightarrow \infty$ , and the optimal strategy is to apply OLS rather than pre-test. This is consistent with the results in Table 1, of course.

In the context of a mini-max regret approach to the choice of  $c^*$ , it seems that little more can be offered by way of a truly general prescription, other than to remark that the correct specification of the initial model is of

paramount importance. Of course, this applies more generally than simply to the choice of  $c$ . Under-fitting the model has other serious implications for the properties of pre-test strategies, as is illustrated by Ohtani (1983) and Mittelhammer (1984), for example.

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Table 1. Optimal Critical Values and Their Significance Levels

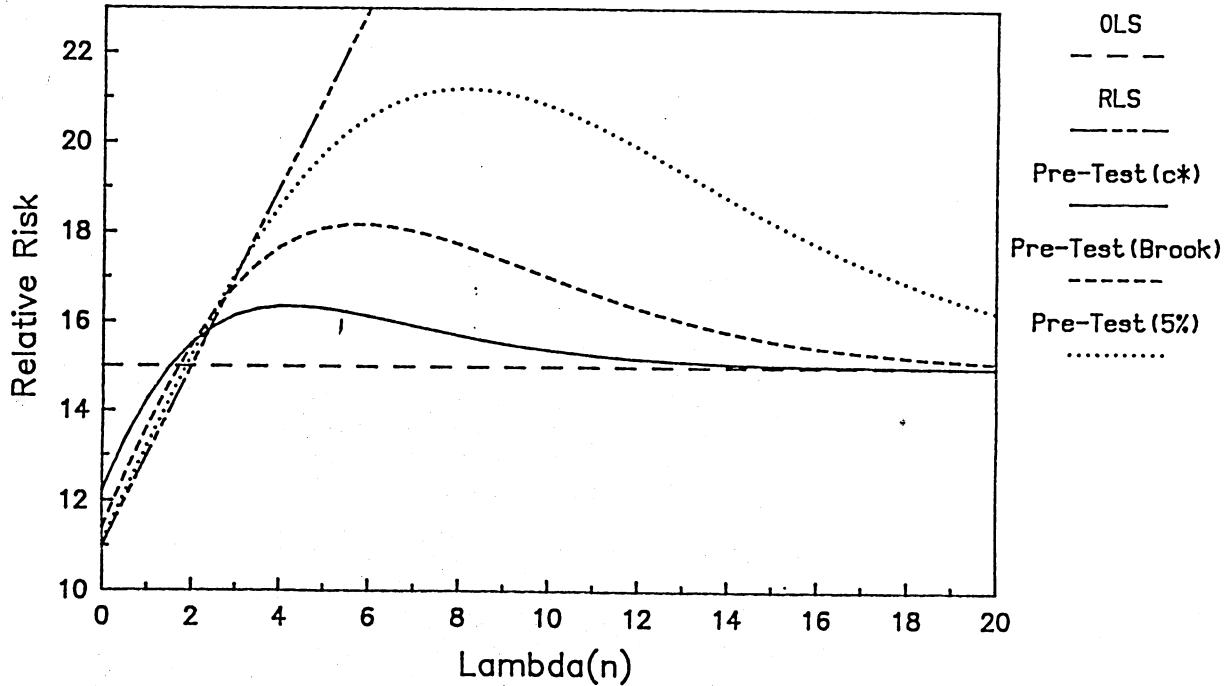
m	$\nu$	$\lambda_d=0$		$\lambda_d=1$		$\lambda_d=5$		$\lambda_d=10$		$\lambda_d=50$		
		c*	$\alpha^*$	c*	$\alpha^*$	c*	$\alpha^*$	c*	$\alpha^*$	c*	$\alpha^*$	
1	2	1.990	0.294	0.970	0.429	0.319	0.629	0.173	0.718	0.037	0.865	
	4	1.940	0.236	1.280	0.321	0.540	0.503	0.317	0.604	0.072	0.802	
	8	1.910	0.204	1.530	0.251	0.840	0.386	0.540	0.483	0.140	0.718	
	16	1.900	0.187	1.690	0.212	1.160	0.297	0.840	0.373	0.260	0.617	
	24	1.890	0.182	1.750	0.198	1.330	0.260	1.028	0.321	0.366	0.551	
	60	1.890	0.174	1.830	0.181	1.620	0.208	1.418	0.238	0.700	0.406	
120	1.890	0.172	1.860	0.175	1.746	0.189	1.621	0.205	1.030	0.312		
2	2	2.090	0.324	1.020	0.495	0.327	0.754	0.176	0.850	0.037	0.964	
	4	2.000	0.250	1.328	0.361	0.558	0.611	0.320	0.743	0.073	0.931	
	8	1.960	0.203	1.558	0.268	0.860	0.459	0.550	0.597	0.141	0.871	
	16	1.930	0.177	1.710	0.212	1.180	0.333	0.850	0.446	0.263	0.772	
	24	1.920	0.168	1.770	0.192	1.350	0.278	1.040	0.369	0.369	0.695	
	60	1.910	0.157	1.850	0.166	1.630	0.204	1.430	0.247	0.710	0.496	
120	1.900	0.154	1.870	0.159	1.760	0.176	1.630	0.200	0.940	0.357		
4	4	2.060	0.251	1.360	0.386	0.564	0.704	0.326	0.848	0.075	0.986	
	8	1.990	0.189	1.583	0.269	0.871	0.521	0.557	0.700	0.144	0.961	
	16	1.960	0.149	1.740	0.190	1.200	0.349	0.865	0.506	0.268	0.894	
	24	1.950	0.135	1.800	0.162	1.378	0.271	1.060	0.398	0.375	0.824	
	60	1.940	0.115	1.880	0.126	1.662	0.171	1.460	0.226	0.727	0.577	
	120	1.940	0.108	1.910	0.113	1.790	0.135	1.660	0.164	1.060	0.379	
8	8	2.020	0.170	1.611	0.258	0.887	0.565	0.567	0.780	0.146	0.993	
	16	1.990	0.115	1.765	0.159	1.220	0.348	0.880	0.553	0.272	0.966	
	24	1.980	0.094	1.826	0.121	1.397	0.248	1.077	0.411	0.381	0.920	
	60	1.970	0.066	1.910	0.075	1.690	0.119	1.480	0.184	0.738	0.658	
	120	1.970	0.056	1.936	0.061	1.820	0.080	1.690	0.108	1.075	0.385	
	16	16	2.007	0.087	1.784	0.129	1.231	0.341	0.887	0.593	0.274	0.993
24		1.999	0.061	1.843	0.085	1.408	0.218	1.087	0.416	0.384	0.974	
60		1.989	0.029	1.923	0.036	1.703	0.071	1.490	0.134	0.745	0.738	
120		1.986	0.019	1.953	0.022	1.832	0.034	1.701	0.055	1.083	0.379	
24		24	2.005	0.048	1.850	0.070	1.413	0.202	1.090	0.417	0.385	0.988
		60	1.994	0.016	1.930	0.021	1.709	0.048	1.495	0.106	0.746	0.783
	120	1.991	0.008	1.958	0.010	1.838	0.017	1.706	0.032	1.086	0.370	
	60	60	2.021	0.004	1.955	0.005	1.731	0.018	1.514	0.055	0.754	0.861
		120	2.011	0.001	1.978	0.001	1.856	0.002	1.723	0.006	1.096	0.332
		120	120	2.010	0.000	1.977	0.000	0.856	0.000	1.723	0.002	1.095

Table 2. Globally Optimal Critical Values and Their Significance Levels

m	$\nu$	$\bar{\lambda}_d$							
		20		50		100		500	
		c**	$\alpha^{**}$	c**	$\alpha^{**}$	c**	$\alpha^{**}$	c**	$\alpha^{**}$
4	16	0.858	0.510	0.471	0.756	0.262	0.898	0.056	0.944
8	16	0.893	0.544	0.498	0.840	0.275	0.965	0.059	1.000
4	60	1.417	0.239	1.049	0.390	0.732	0.574	0.202	0.936
8	60	1.450	0.195	1.077	0.391	0.766	0.633	0.210	0.988



FIGURE 1  
 RELATIVE RISK FUNCTIONS  
 ( $\Lambda(d)=5$ ;  $m=4$ ;  $v=16$ ;  $k=5$ )



NOTE : Brook's  $c=1.96$ ; 5%  $c=3.01$ ;  $c^*=1.2$

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