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Post-Sample Prediction Tests for Generalized Method of Moment Estimators

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1. Introduction

Many models are estimated by method of moments. Hansen (1982) proposed the generalized method of moments GMM estimator for use in macroeconometric work, and many applications of this idea have been made, e.g., Hansen and Singleton (1982), Miron (1986), Dunn and Singleton (1986). Moreover, since the maximum likelihood estimator (MLE) is a sub-set of the GMM class in which the scores constitute the first order conditions the principle extends very widely.

Formally let $g_i(\theta)$ $(i = 1, \ldots, N)$ constitute a m x 1 vector of first order conditions for N observations, θ be a (K x 1) vector of unknown parameters with true value θ_o and $E(g_i(\theta_o))$ equals zero. Hansen's idea was to choose the $\hat{\theta}$ that minimizes $[N^{-1} \sum_{i=1}^{N} g_i(\hat{\theta})]' W [N^{-1} \sum_{i=1}^{N} g_i(\hat{\theta})]$, where W is a pre-defined (m x m) weighting matrix. If $m = K$, $W = I_m$ and $\hat{\theta}$ solves $N^{-1} \sum_{i=1}^{N} g_i(\hat{\theta}) = 0$. When $m > K$ a test for the adequacy of the estimated model is available by seeing how close $N^{-1} \sum g_i(\hat{\theta})$ is to zero, and this constitutes the identifiability test statistic (J) given by Hansen and which appears with most GMM estimates.

The J statistic has been fairly effective in detecting poor models and initially few estimated models survived its application, but this seems to be now changing as authors adopt richer specifications and are more careful in their treatment of the data. Consequently, it may be necessary to augment the J statistic with some other measures of model adequacy. If one examines popular GMM estimators such as OLS, 2SLS, MLE there are a wide range of tests to assess potential departures from the conventions underlying the construction of those estimators, e.g., one has tests for serial correlation, heteroskedasticity, normality etc. But most of these are irrelevant in a pure GMM framework since it is only the first order conditions which need to be satisfied.

It is instructive to observe however that in the most popular GMM estimator of them all--OLS--researchers have found it very useful to examine the predictive ability of a model as a check on its performance. Indeed it would not be a distortion to say that many researchers see predictive success as a critical check on the adequacy of a model. A simple predictive test for OLS estimated models ($y_i = x_i\beta + u_i$) would be to examine the average prediction error over a post-sample prediction period of N+1, ..., N+n, i.e., $n^{-1} \sum_{i=N+1}^{N+n} \hat{u}_i$. A generalization would be to consider the vector $n^{-1} \sum_{i=N+1}^{N+n} x_i'\hat{u}_i$; if x_i contains unity as a regressor this is a super-set of the simple criterion.

Cast into the GMM framework it is $n^{-1} \sum_{i=N+1}^{N+n} \hat{g}_i$ which is being examined, where $\hat{g}_i$ is g_i evaluated with the GMM estimate $\hat{\theta}$, an equivalence which suggests that the average values of the orthogonality conditions in the post-sample prediction period are likely to be a useful diagnostic test.[1] The idea is particularly appealing in situations when n and N can both be made large, since it is likely that the greater n is the more stringent the test of the model will be. This situation occurs mostly in work with financial data and with data on "individuals." In the last category there has been surprisingly little work utilizing predictive ability as a basis for judging estimated Tobit, Logit, Probit, duration models etc. Anderson (1987) is an exception, using the likelihood ratio test for structural change in a model when it is estimated over $i = 1, \ldots, N$ and then $i = 1, \ldots, N+1$. Certain problems occurred when implementing this idea for the Tobit model that the procedure advocated here of focusing upon $\hat{\tau} = n^{-1} \sum_{i=N+1}^{N+n} \hat{g}_i$ does not suffer from. Moreover, examination of the individual elements of $\hat{\tau}$ can be informative about the source of the predictive failure, and this information is not readily available from a likelihood ratio test.

The next section briefly describes the derivation of the asymptotic variance of $n^{1/2}\hat{\tau}$ under the null hypothesis that the model is an adequate one and that $n = kN$, so that both the prediction and sample period become infinitely large together. Such a treatment of prediction seems to accord better with the way prediction tests are frequently used as diagnostic devices. Moreover, without this assumption it would prove very hard to derive

[1]In fact experience from the regression framework indicates that tests based on comparing $n^{-1} \sum_{i=N+1}^{N+n} \hat{u}_i^2$ to $N^{-1} \sum_{i=1}^{N} \hat{u}_i^2$ may be more powerful and one might follow up this idea in the GMM context as well.

the distribution of $\hat{\tau}$, unless very stringent assumptions are made about g_i. Section 3 of the paper applies the test to a range of well known empirical studies.

2. The Asymptotic Distribution of $n^{1/2}\hat{\tau}$

Three assumptions will be made, application of which immediately gives the limiting distribution of $n^{1/2}\hat{\tau}$.

(i) g_i is a stationary and ergodic process with $E(g_i) = 0$ when the model is adequate.

(ii) g_i exhibits sufficiently weak dependence and enough moments exist so that when the model is adequate

$$\begin{bmatrix} n^{-1/2} \sum_{i=N+1}^{N+n} g_i \\ N^{-1/2} \sum_{i=1}^{N} g_i \end{bmatrix} \xrightarrow{d} \mathcal{N}\left[0, \begin{bmatrix} V_g & 0 \\ 0 & V_g \end{bmatrix}\right]$$

as $N \to \infty$, where $V_g = E\left[(N^{-1} \sum_{i=1}^{N} g_i)(N^{-1} \sum_{i=1}^{N} g_i)'\right]$.

(iii) The GMM estimator $\hat{\theta}$ when the model is adequate has the limiting property $N^{1/2}(\hat{\theta} - \theta_o) = A^{-1}N^{-1/2}\sum_{i=1}^{N} g_i + o_p(1)$ so that

$$N^{1/2}(\hat{\theta} - \theta_o) \xrightarrow{d} \mathcal{N}(0, A^{-1}V_g(A')^{-1}).$$

where $A = (\operatorname{plim}_{N\to\infty} N^{-1} \sum_{i=1}^{N} \frac{\partial g(\theta_o)}{\partial \theta}) W = GW$ and G is a $K \times m$ matrix of derivatives.

Assumption (i) is employed by Hansen while (iii) is his result on the asymptotic distribution of the GMM estimator. (ii) holds under a wide range of processes. The diagonality of the covariance matrix is obvious for the case when the g_i are independently distributed, but is also true when the g_i follow stationary ARMA processes and obey various mixing conditions. To appreciate why this is so let g_i be an AR(1) $g_i = \rho g_{i-1} + V_i$. with V_i being iid(0,1). Then the covariance is

$$E\left[(n^{-1/2} \sum_{i=N+1}^{N+n} g_i)(N^{-1/2} \sum_{i=1}^{N} g_i)\right] = k^{-1/2}N^{-1}E\left[g_{N+1}(\sum_1^N g_i) + g_{N+2}(\sum_1^N g_i) + \ldots\right]$$

$$= k^{-1/2}N^{-1}\left[\sum_{j=1}^{N} \rho^j + \sum_{j=2}^{N} \rho^j + \ldots\right] = k^{-1/2}N^{-1}[(1-\rho)^{-2}\rho + o(1)]$$

as $N \to \infty$, which tends to zero asymptotically.[2]

Now the test statistic is $\hat{\tau} = n^{-1} \sum_{i=N+1}^{N+n} \hat{g}_i$ and using the delta method to get the asymptotic distribution yields

$$n^{1/2}\hat{\tau} = \left[n^{-1/2} \sum_{i=N+1}^{N+n} g_i\right] + \bar{G}'(\hat{\theta} - \theta_o) + o_p(1) \qquad (1)$$

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[2]Ghysels and Hall (1987) give a proof of (ii) imposing mixing conditions upon g_i. Note that what is important is to prove joint normality of each of the sample means and this requires an array theorem for mixing processes such as Gallant (1987, p. 519). Our work was done independently of Hall and Ghysels and we have only recently become aware of it.

where $\overline{G}$ is the prediction period analogue of G. Using (iii)

$$n^{1/2}\hat{\tau} = \left[n^{-1/2}\sum_{i=N+1}^{N+n} g_i\right] + n^{1/2}N^{-1/2}\overline{G}'A^{-1}N^{-1/2}\sum_{i=1}^{N} g_i \qquad (2)$$

and then (ii) gives

$$n^{1/2}\hat{\tau} \xrightarrow{d} \mathcal{N}(0, V_g + k\,\overline{G}'A^{-1}V_g(A')^{-1}\overline{G}) \qquad (3)$$

which is what we generally use to construct a $\chi^2(m)$ statistic in Section 3.[3] However, when $\hat{\theta}$ is an MLE, as is the case with most Tobit, Logit, Probit models etc., the g_i are the scores with respect to θ, $m = K$, $W = I_m$, and $\overline{G} = G = A = V_g^{-1}$ when the model is adequate. Hence for these estimated models

$$n^{1/2}\hat{\tau} \xrightarrow{d} \mathcal{N}(0, V_g(1 + k)) \qquad (4)$$

and this is what we utilize for models estimated by ML in the next section.

3. Applications

The focus of our applications is the models estimated by Hansen and Singleton (1982) (HS), Dunn and Singleton (1986) (DS), Miron (1986) and Fair (1977). In the first two papers the authors apply GMM to sample orthogonality conditions to estimate preference parameters embodied in an asset pricing framework. Miron examines the effect of nonseasonal data in a similar "life-cycle-permanent income" model of consumption. Miron's preference parameters are estimated by applying two-stage nonlinear least squares to a system of stochastic Euler equations. Fair fits a Tobit model by M.L. to data on extramarital affairs. While these models have generated substantial

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[3] GAUSS Code that may be added to Runkle and Leonard's GMM program to compute the test statistic is available from the first author.

interest in both econometric and economic issues, our interest is exclusively in the relative performance of the τ and J tests.

We examine a subset of the models estimated in each of the original papers. For HS we concentrate on their Table 1: NLAG=1 estimates. The only DS models considered are the simple models that rely on one-month returns (TBILL) and the three month return from rolling over one-month Treasury Bills (ROLL1). These results are displayed in Table 1 of their paper. Our concern with Miron is the estimation of his equation 18--results presented in Table 2 of his paper. In Fair's case the data used was the 601 observation "Psychology Today" set, and the model estimated is described in Table 3 of his paper.

Data

The data used in our study was obtained from the authors.[4] Details on original sources can be obtained from their papers. The HS/DS data was updated through 1982 using CRSP monthly data on stock returns, Fama's TBILL data, and unrevised (base year 1972) consumption data. This updated data matched well with the existing 1959-1978 sample. To ensure the accuracy of the "in-sample" data and estimation techniques used in this paper, we matched the parameter estimates published in these three papers prior to calculating the specification tests discussed below.

Results

The HS sample period J-tests in the upper half of Table 1 are

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[4]We are grateful to these authors for providing us with their data.

essentially the same as those associated with the NLAG-1 models presented in Table I of the Errata (1984) for their paper.[5] Models estimated with equally weighted equity returns (EWR) are rejected by the J-test of overidentifying restrictions. In contrast, τ based on the 1979-1982 prediction period is well within a conventional acceptance region. The picture changes substantially when we reduce the estimation sample to 1959-1968. In this case no .05 level J-test rejections appear. However, the τ statistic rejects at the .050 level for VWR models and even more convincingly in the EWR models. "t-statistics" associated with individual orthogonality conditions embodied in τ indicate that the lagged equity yield and the constant term (essentially the sum of the Euler equation over the prediction period) prove to be responsible for the rejections.[6]

In an exercise that is preliminary to examination of a more sophisticated model, DS investigate the HS formulation using various real Treasury Bill returns in place of equity returns. In Table 2 J-test rejections are observed in the 1959-1978 sample for each return. More severe rejections are revealed by τ in the 1979-1982 prediction period. Again, significant nonzero Euler equation sums--perhaps associated with the uncharacteristically high real t-bill returns in 1981 and 1982--appear to have precipitated the rejections.[7] In the 1959-1968 sample the J-test rejects the

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[5]Minor discrepancies appear in the EWR results since we used revised CRSP data for EWR.

[6]These ranged from -1.79 to -1.97 in the four models. Detailed estimates of individual orthogonality conditions and t-tests are available on request.

[7]Returns in late 1981 and 1982 averaged two to three times higher than in earlier periods. In these tests the t-values ranged from -1.92 to -3.20. A τ test over the 1979-1980 prediction period only did not result in rejection.

model for the TBILL1 return but not the ROLL1 return. Alternatively, the 1969-1978 τ test rejections are definitive for both returns and associated t-tests reveal that all instruments play a significant role in the rejections.

Results from our re-estimation of Miron's equation 18 are presented in Table 3. Due to the nonstationarity of instruments used in Miron's study, the reported statistics may not have Chi-squared distributions. Nevertheless, we report χ^2 marginal significance levels to facilitate comparison with Miron. The J-test results differ from those reported by Miron for several reasons. First we used 1947-1978 (or 1960-1978) while the Miron sample extended to 1982. Second, Miron informed us that his Table II J-test statistics contained a minor error. Third, Miron calculated J-tests using the NR^2 "residual" based calculation that assumes conditional homoskedasticity, whereas our J-test calculation accounts for possible heteroskedasticity. Still, our "in sample" findings compare favorably with those of Miron--with only two severe rejections in the 1947-1978 period. The τ evidence is slightly more negative in the 1979-1982 prediction period with three of the six categories rejecting at the .025 level. Longer prediction periods in the second half of Table 3 provide more damning evidence against the model--with five of six categories exhibiting highly significant values for $\hat{\tau}$. This contrasts with the fact that there is little evidence against the model conveyed by the shorter sample J-tests. Individual t-statistics associated with each τ reveal that the 1979-1982 rejections may be ascribed to the trend and seasonal dummies, whereas the more severe rejections in the longer "out-sample" periods are generally associated with uniform rejection of all ten orthogonality conditions.[8]

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[8]Detailed estimates of individual orthogonality conditions and t-tests are available on request.

Fair's equation involves the explanation of the number of times extra-marital intercourse took place as a function of nine explanatory variables. The 601 observations were partitioned into 541 for the sample and 60 for the prediction period, making the latter almost 10% of the total data available. These 60 observations were randomly selected. Fair fitted a Tobit model to the data by M.L. so that the g_i are the scores and the overall test statistic was computed using the simplification in (4), giving a value of 61.48. Individual "χ^2-statistics" were 87.75 (constant), 3.08 (sex), 28.29 (age), 2.93 (number of years married), 4.28 (children), 6.60 (religious attachment), 127.5 (education level), 20.08 (occupation) and 19.85 (error variance). Thus there is obvious evidence of inadequacy in the basic model formulation.

4. Conclusion

In this paper we have argued that the magnitude of the post-sample orthogonality conditions associated with GMM estimation can be a useful *adjunct* to standard tests for mis-specification. The test statistic is easily computed and in the examples of Section 3 we found that it frequently gave opposite indications to the identifiability test statistic. Hence there is a strong case that both statistics should be computed routinely. We also argued that the predictive test statistics could be particularly useful for models using individual data, since then the "prediction sample" can be made quite large in absolute terms and yet the sample chosen for estimation will still be sizeable.

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Table 1

J-test and τ test statistics for NLAG=1 estimates in Table 1 of Hansen/Singleton (1982, 1984)

Category	Sample		J-test with df:1		τ with df:3	
	J-test	τ	Value	Prob	Value	Prob
NDS EWR	59-78	79-82	4.89	.973	.38	.056
NDS VWR	59-78	79-82	1.06	.697	.37	.054
ND EWR	59-78	79-82	6.85	.991	.64	.113
ND VWR	59-78	79-82	1.44	.770	.71	.129
NDS EWR	59-68	69-78	2.26	.867	17.76	.999
NDS VWR	59-68	69-78	.93	.665	7.90	.952
ND EWR	59-68	69-78	2.57	.891	20.06	.999
ND VWR	59-68	69-78	1.15	.716	7.07	.930

All estimates are obtained using a GMM program written by Runkle and Leonard. Data was obtained directly from the authors and is described in the original paper. Original sample 59-78 results based on VWR match those presented in Errata (1984) Table 1. EWR results differ slightly from H/S since we only had access to revised CRSP data on EWR. The instruments used to estimate the models and form the test statistics include a constant, a one period lag of the percent growth in real per capita consumption, and a one period lag of equity returns. Two parameters were estimated in each model. t-statistics for individual orthogonality conditions used in τ along with all parameter estimates are available on request.

Table 2

J-test and τ test statistics for models estimated
by Dunn and Singleton (1986)
(one month T-bills in D/S Table 1)

Category	Sample		J-test with df:3		τ with df:5	
	J-test	τ	Value	Prob	Value	Prob
TBILL1	59-78	79-82	15.49	.999	111.94	.999
ROLL1	59-78	79-82	10.83	.987	56.06	.999
TBILL1	59-68	69-78	11.65	.991	189.20	.989
ROLL1	59-68	69-78	5.38	.854	122.60	.989

All estimates are obtained using a GMM program written by Runkle and Leonard. Data was obtained directly from the authors and described in the original paper. Parameter estimates and test statistics essentially match those presented in Table 1 of DS. The ROLL1 estimates account for a second order moving average process precipitated by the three month return used in the model. The instruments used to estimate the models and to form the test statistics include a constant, two lagged values of the real per capita growth rate of non-durable plus service consumption and two lagged values of real t-bill returns. Two parameters were estimated in each model. t-statistics for individual orthogonality conditions used in τ along with all parameter estimates are available on request.

Table 3

J-test and τ test statistics for equation 18 (Miron, JPE 1986)

Category	Sample		J-test with df:4		τ with df:10	
	J-test	τ	Value	Prob	Value	Prob
FOOD	48:1-78:4	79:1-82:4	3.99	.592	7.80	.351
GO	48:1-78:4	79:1-82:4	6.43	.830	2.38	.007
FOC	60:1-78:4	79:1-82:4	.12	.002	30.24	.999
HOUS	48:1-78:4	79:1-82:4	67.10	.999	23.42	.991
EG	60:1-78:4	79:1-82:4	2.41	.339	34.02	.999
TRANS	48:1-78:4	79:1-82:4	20.12	.999	5.71	.161
FOOD	48:1-64:2	64:3-82:4	.38	.016	15.17	.874
GO	48:1-64:2	64:3-82:4	.35	.014	40.93	.999
FOC	60:1-71:2	71:3-82:4	1.16	.115	199.53	.999
HOUS	48:1-64:2	64:3-82:4	3.45	.514	461.79	.999
EG	60:1-71:2	71:3-82:4	19.78	.999	194.08	.999
TRANS	48:1-64:2	64:3-82:4	5.13	.726	32.58	.999

All estimates were obtained from a base program written for TSP by Miron. Data was obtained directly from Miron and is described in the original paper. J-tests differ slightly from results in Miron (1986) Table 2 for reasons discussed in the text. Test statistics may only approximate a Chi-squared distribution due to the nonstationarity of the data used by Miron. Instruments used in the estimation and construction of tests include a constant, trend, trend squared, two lags of real interest rates, two lags of the growth in real per capita consumption, and three seasonal dummies. Contrary to Miron's Table 2, <u>six</u> parameters are estimated in equation (18).

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