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STOCHASTIC AND DETERMINISTIC TRENDS IN FINNISH MACROECONOMIC TIME SERIES*

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The appropriateness of the Dickey-Fuller unit root test is studied using two alternative unit root test models. The segmented trend model is strongly supported and the second-order trend model is favoured for some series. The sample set consists of observations of nine basic macroeconomic time series describing the fundamentals of the Finnish economy between 1860 and 1989. The results clearly show that care is required in interpreting unit-root tests since failure to reject does not entail that the null is true. Structural breaks in the data generating process, in this case wars starting in 1917 and 1939, support models of the deterministic trends class. However, it is argued that the univariate testing procedures laid down in the unit root literature do not provide information to macroeconomic controversies. (JEL C22)

1. Introduction

Until recently, macroeconomic time series were widely viewed as being *trend-stationary* (TS), comprising stationary fluctuations around deterministic linear or exponential trends. Random innovations only had a temporary influence on the historical trajectory of the series. The work of Fuller (1976) and Dickey & Fuller (1981) on the econometrics of integrated time series, however, made testing for the presence of unit autoregressive roots straightforward, and Nelson & Plosser (1982) popularized it in the field of macroeconomics. If a particular series has a unit root (i.e. *difference stationarity* (DS) with, possibly, a non-zero drift, then the trend is stochastic and involves the accumulation of random innovations, each of which has an enduring effect on the future trajectory of the time series.

Nelson and Plosser used the Dickey-Fuller test procedure to distinguish between these two hypotheses on a selected group of major U.S. macroeconomic series. They found that all but one of the series tested had a unit root (the exception being unemployment), and they concluded that the non-stationary of the macro time series is in general stochastic. This result has a potential implication for economic theory. Transitory shocks must be relatively unimportant in determining the movement of this series, while permanent shocks must dominate in the long run.

On statistical grounds, the approach used by Nelson and Plosser is not without problems. It is argued that, in finite samples, it is almost impossible to test effectively the DS vs. TS hypothesis. Integration tests have low

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power against relevant TS alternatives (DeJong, Nankervis, Savin & Whiteman 1988 and 1992, Cochrane 1991). The Dickey-Fuller tests rely on the innovation process being white noise, and so these tests are not appropriate if the innovations are moving averages (Schwert 1987, Phillips & Perron 1988). When the series include major structural breaks in the level and slope of trend, the tests give misleading inferences (Perron 1989, 1990, Rappoport & Reichlin 1989, Hendry & Neale 1991). If the underlying data generating process is actually random walk *with drift* rather than pure random walk, the power of the standard unit root tests on the deterministic components is low, and the use of Dickey-Fuller tables can be misleading in small samples (Hylleberg & Mizon 1989, Haldrup & Hylleberg 1989).[1]

This paper describes two alternative models for testing the DS vs. TS alternative on a long run macroeconomic series of the Finnish economy. The Dickey-Fuller approach is used augmented with shifts and breaks in the mean of the series (i.e. in the trend) modelled as proposed by Perron (1989, 1990) and Rappoport & Reichlin (1989). The second alternative is the test suggested by Ouliaris, Park & Phillips (1989). They have developed tests for detecting a unit root with higher-order trends under the maintained hypothesis. An important feature of these tests is that they include weak dependence and heterogeneity in the data generating process. Unlike conventional unit root tests, they allow explicitly for deterministic non-stationarity of an arbitrary, but known, form. The focus in this paper is on the linear second-order polynomial trend model.

The tests and series used are introduced in Section II. Section III reports the main findings. The results are provocative. First, the segmented trend model specification is more appropriate than the second order and difference stationarity alternatives. Second, for some series, the 2nd order trend model fits better than the stochastic trend model. Section IV presents a summary of the paper and includes with comments about the unit root testing of macroeconomic series.

2. *Data series and test procedures*

2.1 *Finnish macro series: annual observations from 1860–1989*

The data set consists of eight macroeconomic time series with annual observations from 1860–1989, and one from 1900–1989 (lnIEi00). The series tested are as follows:

lnGNP26 = log of GNP at 1926 prices,
lnPCi26 = log of private consumption index at 1926 prices,
lnEMP = log of employment,
lnPROD26 = ln(GNP26/EMP),
lnEXP26 = log of exports at 1926 prices,
lnINV26 = log of gross investments in machinery and production equipment at 1926 prices,
lnGNPci26 = log of GNP per capita index at 1926 prices,
lnCLi26 = log of cost of living index at 1926 prices,
lnIEi00 = log of the average total earnings index of industrial workers at 1900 prices.

The series (except for the years 1986–1989) and the lnIEi00 series) were taken from the final report of the economic growth research project of the Bank of Finland (Hjerppe 1989). The quality of data is high. The lnIEi00 series was taken from the publication of the main Finnish labour market organizations (STK–SAK, 1988). The observations for the years 1986–1989 are from the Bulletin of Statistics 1990:I (Central Statistical Office of Finland).

2.2 *Test models*

Three different test models were used under the null hypothesis of a unit root.

Model A:

$$y_t = \mu + \beta t + \alpha y_{t-1} + \sum_{1}^{4} c\Delta y_{t-k} + \varepsilon_{1t},$$

Model B:

$$y_t = \mu + \beta t + \delta DTB1917 + \alpha y_{t-1} + \varphi_1 DU1917 + \varphi_2 DU1939 + \gamma DT1917 + \sum_{1}^{4} c\Delta y_{t-k} + \varepsilon_{2t}.$$

[1] *The aim is not to give a review of the current state of the art in unit root research. Many excellent reviews already exist (see for example Pagan & Wickens 1989, Bell & Miller 1986, Dolado, Jenkinson & Sosvilla-Rivero 1990, Diebold & Nerlove 1990).*

Model C:

$$y_t = \mu + \beta_1 t + \beta_2 t^2 + \alpha y_{t-1} + \varepsilon_{3t}$$

Model A) is the familiar augmented Dickey-Fuller (ADF) test. The null hypothesis is $\alpha = 1$ against the alternative $\alpha < 1$. An obvious test statistic is the usual 't-ratio' of the $(\alpha - 1)$ estimate. However, Dickey and Fuller show that this statistic does not have a Student's t-distribution. Thus, the critical 't-values' for the parameters in model A are non-standard and were tabulated using Monte Carlo methods, see Dickey and Fuller (1976, 1981) and Guilkey & Schmidt (1989). The test is valid even if the innovations series are taken to be the ARMA(p, q) type with the unknown orders p and q. This allows the $\{y_t\}$ series to represent quite a general process. This is approximated with an AR(4) process.

Model B is the structural break model introduced by Perron (1989). It is similar to the ADF-model but, under the null hypothesis $\alpha = 1$, also contains exogenous structural breaks at time points 1917 and 1939. Variables DU1917 and DU1939 are dummy-variables taking value 1 after years 1917 and 1939. Otherwise they are 0. Variable DT1917 is a trend variable starting from 1917. Variable DTB1917 takes value 1 in 1917. Otherwise it is 0. Thus, in 1917 the intercept and the slope of the trend function change. In 1939 only a permanent change in the intercept of the trend function takes place. The structural change time points 1917 and 1939 are assumed to be *given and known*. These dates are evidently exogenous in view of the importance of these years in Finnish history. In 1917, the process leading to independence and the Civil War started. In 1939, the war (Winter War) started against the Soviet Union, leading to wars against the Soviet Union in 1941 – 44 and Nazi Germany in 1944 – 1945. These incidents had major effects on the capital stock and human resources of the Finnish economy.

Thus, under the alternative hypothesis of »trend stationary» process, it is to be expected that $\alpha < 1$; β, φ_1, φ_2 and $\gamma \neq 0$. However, Perron (1989) and Rappoport & Reichlin (1989) (see also Hendry & Neale 1991) showed that ADF-tests are not consistent against trend stationary alternatives when the trend function contains a shift in the slope. The shifts in the intercept of the trend reduce the power of the DF-tests for α. The critical values of the DF-test for α have to be bigger in absolute value when structural breaks are present in the trend part of the model, and these values depend on the break times T_{Bi}. Perron uses Monte Carlo methods to tabulate critical values for model type B) with different values of $\lambda = T_B/T$, where T_B is the time point of the trend break.

Model B does not fulfill the requirements of Perron's approach, because he uses *one* break time point, whereas model B has two (i.e. 1917 and 1939). However, it is known that the critical values have to be about 20 % higher in the case of two mean process shifts compared with the one-shift case (see Table 1. in Rappoport & Reichlin 1989). Thus, if the estimated absolute t value of α is much bigger than the 1 % critical level in Perron's Tables, the DS-alternative can safely be rejected.

Model C is the 2nd-order deterministic time polynomial model suggested by Ouliaris, Park and Phillips (1989). It is quite general because it allows weak dependence and hetogeneity in the error process (see also Phillips 1987). The null hypotheses are $\alpha = 1$ and $\beta_2 = 0$. If the null hypotheses are rejected, the growth processes of the relevant macroeconomic series show increasing or decreasing growth trends.

The test procedures for the model alternative are as follows:

$$S_2(\hat{\alpha}) = (\hat{\sigma}/\hat{\omega})t_{\hat{\alpha}} - T(\hat{\omega}^2 - \hat{\sigma}^2)/(2\hat{\omega}s_0), \text{ and}$$

$$G_2(\hat{\alpha}, \hat{\beta}_2) = (\hat{\sigma}^2/\hat{\omega}^2)F_W + T(\hat{\omega}^2 - \hat{\sigma}^2)^2/(4\hat{\omega}^2 s_0^2) - T(\hat{\alpha} - 1)(1 - (\hat{\sigma}/\hat{\omega})^2),$$

where

$$\hat{\sigma}^2 = \mathrm{var}(\hat{\varepsilon}_{3t}),\ \hat{\omega}^2 = 1/T \sum_1^T \hat{\varepsilon}_{3t}^2 + 2/T \sum_{k=1}^{l} w_l(k) \sum_{t=k+1}^{T} \hat{\varepsilon}_{3t}\hat{\varepsilon}_{3t-k},$$

$$w_l(k) = 1 - k/(l+1),$$

s_0^2 = residual sum of squares from the regression of y_{t-1} on 1, t and t^2,

l = lag window length, $l = 4$.

$t_{\hat{\alpha}}$ = t-value of $\hat{\alpha}$ in the above OLS-regression (C), and

F_W = Wald-test statistics for restrictions $\alpha = 1$, $\beta_2 = 0$.

Ouliaris, Park and Phillips use Monte Carlo methods to tabulate the critical values of tests

$S_2(\hat{\alpha})$ and $G_2(\hat{\alpha}, \hat{\beta}_2)$. The Dickey-Fuller test statistics are not appropriate because, in model C), unit root testing takes place under maintained trends.

3. *The results*

The results from the standard ADF(4)-testing are depicted in Table 1.

All except the employment series lnEMP have a unit root. The unit root alternative is rejected for series lnEXP26 and lnIEi00 at the 20 % level. Thus, as Nelson & Plosser (1982) suggest, it might be argued that the long run macroeconomic series of the Finnish economy are stochastic trends. However, such an inference is not made, since the testing procedure may not be valid or its power may be low.

Table 1. $y_t = \mu + \beta t + \alpha y_{t-1} + \sum_1^4 c\Delta y_{t-k} + \varepsilon_{1t}$.

	T	$\hat{\mu}$	$t_{\hat{\mu}}$	$\hat{\beta}$	$t_{\hat{\beta}}$	$\hat{\alpha}$	$t_{\hat{\alpha}}$	$S(\hat{\varepsilon}_1)$
lnGNP26	125	.3021	1.291	.0010	1.512	.9241	−1.216	.0394
lnEMP	125	3.1362	3.885^{3}	.0026	3.792^{3}	.7761	−3.871^{2}	.0209
lnPROD26	125	−.0061	−.202	.0005	1.589	.9853	−.840	.0336
lnEXP26	125	2.0737	3.036^{1}	.0065	2.891^{2}	.8091	−2.964	.2621
lnINV26	125	1.0384	2.574	.0039	2.559^{3}	.9036	−2.481	.1252
lnGNPci26	125	−.0215	−.704	.0007	1.496	.9776	−1.020	.0458
lnPCi26	125	.5003	2.630	.0021	2.016	.9316	−1.904	.0632
lnCLi26	125	.0198	1.307	.0035	2.772^{2}	.9572	−2.504	.1329
lnIEi00	85	.1196	3.309^{2}	.0039	2.969^{2}	.8674	−2.848	.0705

NOTE: $^{3, 2}$and ^{1} denote statistical significance at the 1 %, 5 % and 10 % levels respectively (see Table 8.5.2 in Fuller 1976, Table 1. in Guilkey & Schmidt (1989) and Tables II and III in Dickey & Fuller 1981).

Table 2. $y_t = \mu + \beta t + \delta DTB1917 + \alpha y_{t-1} + \varphi_1 DU1917 + \varphi_2 DU1939 + \gamma DT1917 + \sum_1^4 c\Delta y_{t-k} + \varepsilon_{2t}$.

	T	$\hat{\mu}$	$t_{\hat{\mu}}$	$\hat{\beta}$	$t_{\hat{\beta}}$	$\hat{\delta}$	$t_{\hat{\delta}}$
ln GNP26	125	4.2029	4.968	.0087	4.900	−.1158	−3.033
lnEMP	125	5.8337	5.951	.0057	5.698	−.0509	−2.549
lnPROD26	125	−.5271	−5.158	.0038	5.260	−.0781	−2.421
lnEXP26	125	7.1568	6.464	.0221	5.291	−.5016	−2.069
lnINV26	125	2.5983	4.195	.0084	3.639	−.1901	−1.503
lnGNPci26	125	−.0201	−1.358	.0055	4.601	−.1352	−3.066
lnPCi26	125	2.6438	5.259	.0113	5.054	−.0317	−0.493
lnCLi26	125	−.2971	−2.651	.0023	2.012	.4161	3.207
lnIEi00	85	−.4577	−2.326	−.0142	−2.463	−.1108	−1.498

	$\hat{\alpha}$	$t_{\hat{\alpha}}$	$\hat{\varphi}_1$	$t_{\hat{\varphi}_1}$	$\hat{\varphi}_2$	$t_{\hat{\varphi}_2}$	$\hat{\gamma}$	$t_{\hat{\gamma}}$	$S(\hat{\varepsilon}_2)$
lnGNP26	.6577	−4.934^{3}	−.4122	−4.703	−.0731	−3.748	.0057	4.731	.0329
lnEMP	.5575	−5.934^{3}	.0811	4.314	.0356	3.623	−.0015	−4.301	.0193
lnPROD26	.7003	−5.364^{3}	−.4268	−5.295	−.0936	−4.835	.0062	5.315	.0281
lnEXP26	.3372	−6.473^{3}	−1.6212	−5.474	−.6919	−5.043	.0210	5.025	.2291
lnINV26	.6904	−4.357^{3}	−.5253	−3.177	−.0391	−0.763	.0080	3.177	.1197
lnGNPci26	.6044	−4.919^{3}	−.5630	−4.723	−.0763	−3.549	.0084	4.883	.0382
lnPCi26	.6247	−5.204^{3}	−.4165	−4.733	−.1113	−3.450	.0046	4.114	.0569
lnCLi26	.8618	−3.378	−.2153	−1.316	.1007	1.891	.0071	1.998	.1202
lnIEi00	.6918	−3.021	−1.3140	−3.709	.0108	.377	.0236	3.727	.0649

NOTE: $^{3, 2}$ and ^{1} denote statistical significance at the 1 %, 5 % and 10 % levels respectively (see Table VI.B in Perron 1989 with $\lambda_{B2} = .4$ for all the series except lnIEi00, for which the value of .2 is valid).

Table 3. Model $y_t = \mu + \beta_1 t + \beta_2 t^2 + \alpha y_{t-1} + \varepsilon_{3t}$

Series	T	$S_2(\hat\alpha)_{l=1}$	$G_2(\hat\alpha, \hat\beta_2)_{l=1}$	$S_2(\hat\alpha)_{l=4}$	$G_2(\hat\alpha, \hat\beta_2)_{l=4}$
lnGNP26	129	-4.56^3	13.66^1	-4.33^2	12.65
lnEMp	129	-9.35^3	49.62^3	-9.21^3	48.84^3
lnPROD26	129	-3.05	9.52	-2.96	8.98
lnEXP26	129	-8.27^3	18.66^2	-8.75^3	18.22^2
lnINV26	129	-3.36	13.38	-3.71	13.94^1
lnGNPci26	129	-3.83^2	14.65^1	-3.65^1	13.53^1
lnPCi26	129	-2.92	8.57	-2.89	8.41
lnCLi26	129	-1.98	5.77	-2.19	6.39
lnIEi00	89	-2.47	7.21	-2.57	7.65

Series	$S(\hat\varepsilon_3)$	$\hat\mu$	$\hat\beta_1$	$\hat\beta_2$	$\hat\alpha$
lnGNP26	.0412	1.6501	.00148	.0000016	.8669
lnEMP	.0207	2.8872	.00235	−.0000045	.7946
lnPROD26	.0333	−.1891	.00227	.0000161	.8869
lnEXP26	.2593	2.5761	.00828	.0000413	.7567
lnINV26	.1288	1.5951	.00602	.0000222	.8477
lnGNPci26	.0466	.1410	.00363	.0000233	.8382
lnPCi26	.0615	.7515	.00325	.0000093	.8931
lnCLi26	.1439	.0470	.00329	−.0000016	.9638
lnIEi00	.0722	.1069	.00316	.0000056	.8206

NOTE: [3], [2] and [1] denote statistical significance at the 1 %, 5 % and 10 % levels respectively (see Table III. in Ouliaris, Park & Phillips 1989).

The results from the segmented trend model are depicted in Table 2. All the series except price series lnCLi26 and lnIEi00 are stationary at the 5 % level, and most of them are stationary at the 1 % level irrespective of the 20 % upward correction of Perron's critical values. The structural breaks (i.e. the wars) seem to have only temporarily interrupted the stationary growth trend in the Finnish economy during the years 1860 – 1989, and did not have any permanent long run effects on the growth process. Nonstationary series with structural breaks would give an opposite conclusion.

The test values for the increasing or decreasing trend growth processes (i.e. model alternative C); the second order trend model) are given in Table 3. The test values of the lnEMP, lnEXP26 and lnGNPci26 series are below the 10 % significance level, and stationary around the second-order trend polynomial. Series lnGNP26 and lnINV26 are stationary at the 10 % level, depending on the lag truncation value *l*.

The results shown in Tables 3. and 1. reveal interesting issues for some series. For example, the lnGNP26 series is the unit root series in Table 1., but the unit root is rejected at the 1 % level in Table 3., although the second-order trend term is significant only at the 15 % level. This indicates that the test suggested by Oulirias, Park and Phillips has rather high power compared with the ADF(4)-test mentioned above, although they do report that their tests have a low power when a positive MA(1) term is present in the error process $\{\hat\varepsilon_{3t}\}$ (see Ouliaris, Park & Phillips 1989, p. 25). In the lnGNP26 series, the MA(1) term of process $\{\hat\varepsilon_{3t}\}$ has an ML-estimate of .429.

4. Summary

Although the question of the right model specification is not raised in this paper, it is reasonable to argue that the proposed segmented trend model encompassed at least the (stochastic) trend model. Standard errors of model estimates are lower in the segmented trend alternative compared with the DS alternative. However, the issue is more complex because the above results are somewhat superficial. Observational equivalence exists between the model alternatives in finite samples: trend stationary processes with breaks are almost observationally equivalent to unit root processes with strong mean-reversion and fat-tailed distribution for the error sequence (Perron 1989, Cochrane 1991). Correctly modelled structural breaks »soak up» the large variances and the AR(1) coefficients of the unit root processes.[2] The big and permanent swings in the mean processes are modelled with deterministic breaks and shifting trends. Distinguish between the stationary and difference stationary ARMA model on the basis of a finite data set is almost impossible (Christiano & Eichenbaum 1989). Thus, regime shifts can mimic unit roots in autoregressive time series. But existence of unit roots may be incorrectly assessed without a specific alternative deterministic trend model. The widely used test of Dickey and Fuller thus tends to favour the DS model over the TS model when the true process is a segmented trend.

The fact that the trend changes were modelled as exogenous, and as occurring on known, dates implies that the results described here are conditional, i.e. the fluctuations are transitory because of the presence of a change in trend function in 1917 and 1939. This may cause pretest bias at the significance levels used (see Christiano 1988, Banerjee, Lumsdaine & Stock 1990).

All these arguments indicate that unit root testing is not a free lunch. The testing of the DS and TS alternatives is a complex issue. The results very much depend on the way the problem is tackled. The correct alternative cannot be decided using the data alone but has to be based on the theory or a priori reasoning. However, despite the views of the economic profession in general, keen attachment to the unit root alternative based on *empirical* evidence is alarming. Empirical facts normally do not seem to matter much in the economic profession. As long as unit root testing is a rather fragile procedure and the model equivalence is evident, anything approaching an economic theory cannot be tested effectively using only univariate time series representations.

[2] There has been some progress in using unit root tests in the presence of fat-tailed distributions (see Phillips 1990, Phillips & Loretta 1991, Kim & Schmitd 1989). The results imply that the modified Dickey-Fuller approach suggested by Phillips & Perron (1988) is asymptotically valid except in some exceptional cases (i.e. cases when the characteristic coefficient of stable process is below 1, see Phillips 1990).

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