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International evidence on shock persistence: structural change, nonlinearities and subsample robustness

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Testing for unit roots and the related issue of measuring shock persistence has attracted considerable theoretical and applied econometric interest. The issue of the size of the random walk component raised by Cochrane (*Journal of Political Economy*, **96**, 893–920, 1988) has been extended in the work of other authors. Leung (*Economics Letters*, **40**, 435–44, 1992), in particular, considers the effects of structural breaks on measures of persistence. In this paper we consider new results for the effects of structural change, nonlinearities and subsample robustness on spectral-based measures of persistence illustrating the potential problems via an updated and extended version of the Cogley (*Journal of Political Economy*, **98**, 501–18, 1990) data set. We find that significant structural breaks exist in the majority of the series investigated. Furthermore, measured persistence differs markedly across distinct periods with the assumption of common growth rates over very long periods leading on occasions to potentially erroneous conclusions on the degree of persistence. Nonlinearities measured by the significance of time squared (TSQ), seem important in many cases, particularly post-World War Two (WW2), potentially explaining the apparently high levels of post-WW2 persistence. The paper concludes with a warning on the use of measures of persistence when the data include structural change and/or nonlinearities, highlighting the importance of the correct choice for breakpoints and second sub-sample start points.

I. INTRODUCTION

Testing for unit roots, and the related issue of measuring shock persistence, has attracted considerable theoretical and applied econometric interest. The initial work of Dickey and Fuller (1979) has led to considerations of the theoretical and empirical effects of structural breaks see for example, Perron (1989), Rappoport and Reichlin (1989), Raj (1992) and Zivot and Andrews (1992). Furthermore, the issue of the size of the random walk component raised by Cochrane (1988) has been extended in the work of Cogley (1990), Jaeger (1990), Leung (1992), Mayadunne *et al.* (1995) and Islam (1996). Cogley (1990) considers international evidence on output per capita shock persistence, concluding in favour of an idiosyncratically low measure for the USA consistent with the trend stationary representation identified from the unit root testing framework. Jaeger (1990) and Leung (1992) concentrate on the effects of interpolation on persistence measures following the work of Romer (1989) extending the issues raised by Cochrane (1988). In particular, Cochrane (p. 902) considers the biases in persistence measure results which may arise via interpolation or changes in the growth rates of the series over long periods. Leung (1992) considers the effects of structural breaks on both measures of persistence and also cycle length, leading to a partial resolution of the differing results of Zarnowitz and Moore (1986), De Long and Summers (1988) and Cogley (1990).

In this paper we add to the debate by considering the effects of structural change, nonlinearities and sample period sensitivity, on spectral-based measures of persistence

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illustrating the potential problems via an updated and extended version of the Cogley (1990) data set. In all cases structural change will be identified via a unit root testing framework following the methodologies of Perron (1989) and particularly Zivot and Andrews (1992).

To anticipate, we find that significant structural breaks exist in the majority of the series investigated. Furthermore, and more importantly given the theme of this paper, measured persistence may differ markedly across distinct periods with the assumption of common growth rates over very long periods leading to erroneous conclusions on the degree of persistence. Nonlinearities as measured by the significance of time squared (TSQ) may lead erroneously to high measured levels of persistence. Such conclusions add support to the work of Leung (1992) and reinforce the doubts of Cogley (1990) and Mayadunne *et al.* (1995). In Section II we consider some of the econometric issues of measurement with Section III presenting the empirical examples. Section IV concludes.

II. STATISTICAL METHODOLOGY

One of the first ways to *infer* persistence used tests which categorize univariate series as either Trend Stationary (TS) or Difference Stationary (DS). The class of model most commonly used to describe temporary, i.e. nonpersistent, deviations about a trend is:

$$y_t = a + bt + u_t \tag{1}$$

where y_t is the natural logarithm of the variable of interest, t describes the trend and u_t is a stationary invertible autoregressive moving average (ARMA) process. This process is stationary in levels or trend stationary, (TS). The simplest class of model which captures permanent, i.e. persistent fluctuations is the random walk with drift

$$y_t = \mu_t + y_{t-1} + e_t \tag{2}$$

The random walk is nonstationary in levels, but stationary when differenced, i.e. difference stationary, (DS).

It has become common practice to discriminate between DS and TS processes and hence *infer* persistence or otherwise by using the augmented Dickey–Fuller unit root tests (see Dickey and Fuller, 1981). The usual form of the test treats DS as the null hypothesis and involves estimation of a model like,

$$y_t = \mu + \rho y_{t-1} + \beta t + \sum_{i=1}^{p} \varphi_i \Delta y_{t-i} + \upsilon_t \tag{3}$$

where υ_t is assumed to be serially uncorrelated and the null hypothesis of DS implies H_0: $\rho = 1$, $\beta = 0$. As is now well known under the null hypothesis the usual t ratios are distributed as Dickey–Fuller τ (1981) rather than Student's t. This is the approach pioneered by Nelson and Plosser (1982) and followed on numerous occasions. Perron (1989) demonstrated how breaks in the series can lead to biased results (in favour of DS) and Zivot and Andrews (1992) consider the issue of endogenous versus exogenous breaks. The main differences in the testing procedures proposed by Perron (1989) and Zivot and Andrews (1992) involves the addition of various dummy variables to Equation 3 to capture changes in the intercept and/or time trend and the use of recursive estimation methods, i.e.,

$$y_t = \mu + \rho y_{t-1} + \beta t + \gamma DT_t + \theta DU_t + \sum_{i=1}^{p} \varphi_i \Delta y_{t-i} + u_t \tag{4}$$

where $DU = 1$ if $t > TB$, 0 otherwise and $DT = t$ if $t > TB$ and 0 otherwise and TB refers to the time of the break. In the original Perron formulation, the variables DU and DT were included to capture the possibility of 'crashes' (DU) trend changes (DT) and joint crashes and trend changes (DU and DT). In the empirical section below we consider the possibility of 'jumps' where the coefficient attached to DU could be positive rather than negative as in the case of a 'crash'. Likewise a joint 'jump' and trend change (denoted j and t), would involve the inclusion of both DU and DT in the equation. Tests of the null hypothesis of DS still involve H_0: $\rho = 1$, $\beta = 0$ although critical values are now given in Perron (1989) for exogenous breaks or Zivot and Andrews (1992) for endogenous breaks.

The distinction between a DS and TS process and hence the inference on absolute or no persistence is extreme. Campbell and Mankiw (1987) and Cochrane (1988) emphasize this fact and consider some measure of the *persistence* of shocks, i.e. how much does a one-unit shock to output, y_t (say) affect forecasts into the future? Furthermore, Cochrane (1988) demonstrates how any DS process can be represented as the sum of a stationary and random walk component where the issue of persistence revolves around the *size of the random walk element*. He gives two equivalent formulations of this measure where the second is equal to the spectral density of $\Delta y_t = (1 - L)y_t$ at frequency zero which can be estimated by the Bartlett estimator[1] (see Anderson, 1971). This estimator forms the basis of the Cochrane-type measure of persistence used in this study. However, as demonstrated by Cochrane (1988) the Bartlett estimator will be biased in small samples where the bias can be corrected by multiplying the estimates by $T/(T - k)$ where T is the effective sample size and k the window size.

One of the main advantages of the Cochrane-type measure over parsimonious ARMA representations is that it captures all the effects of a unit root on the behaviour of a series in a finite sample. Furthermore, the size of the random walk element is a continuous choice where series can be more fruitfully categorized via the degree of persistence.

[1] Estimation was undertaken using Microfit 3.0, see Pesaran and Pesaran (1991) and RATS for Windows 4.2, see Doan (1995).

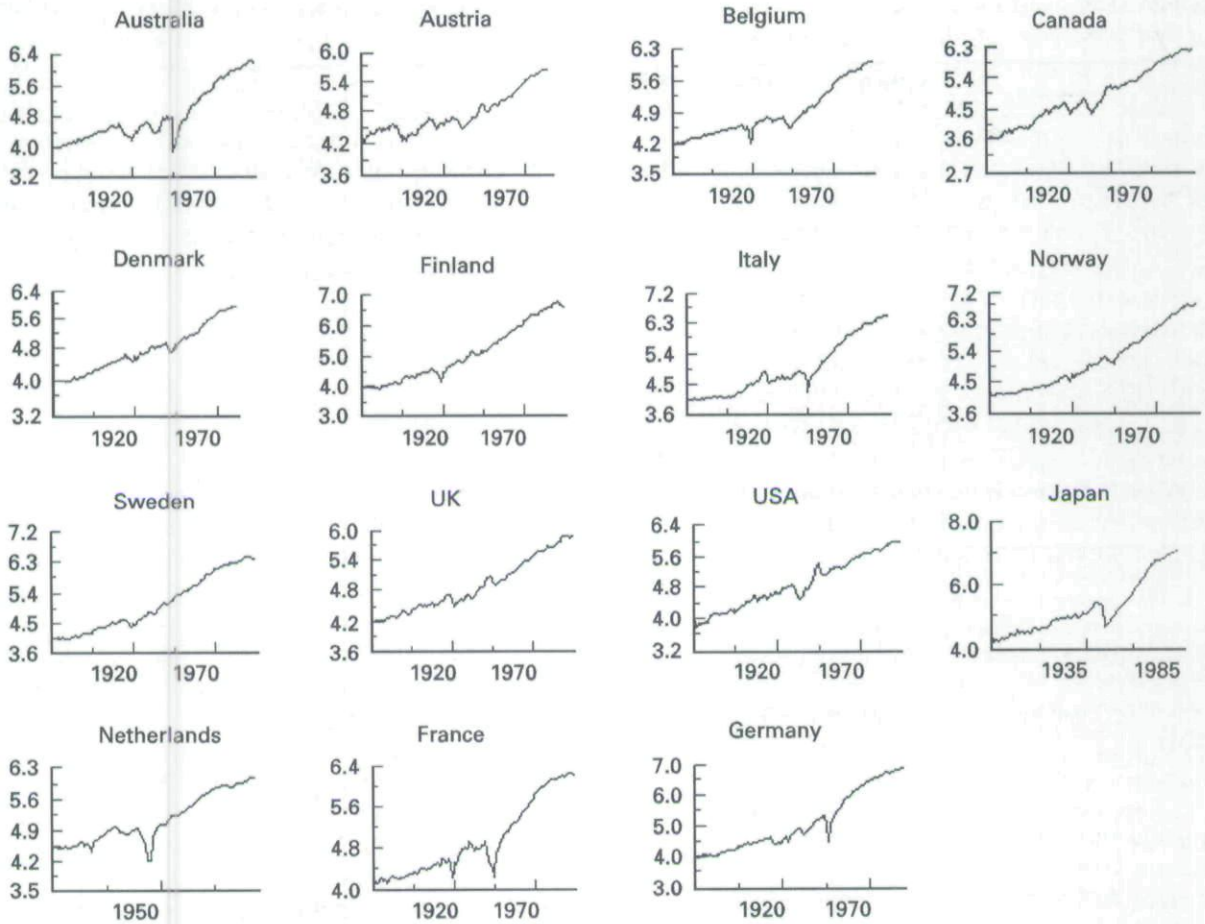

Fig. 1. *Real GDP per capita (logs)*

However, the persistence measure is derived assuming an underlying linear model (without discontinuities). Nonlinearities in the underlying series would invalidate this relationship and the relevance of the Cochrane-type persistence measure.

Consequentially the methodology here has the following steps. Check the order of integration of the data using standard or extended versions of the Augmented Dickey–Fuller test. If the series is TS without breaks measure the degree of persistence for the full sample. If it is TS with breaks use the timing of the breaks to delineate initial subsamples, and measures persistence for these periods. If it is DS after testing for breaks measures persistence for the full sample. Investigate the existence of significant nonlinear elements. Finally check the sensitivity of all results to specific sample periods.

III. RESULTS

Full sample results – all countries

The data used are drawn from Maddison (1991) updated to 1993 with official statistics and relate to the natural logarithm of real GDP per capita. These annual data are available from 1870 except for the Netherlands, France, Germany and Japan. The data, which for Germany are terminated in 1989 to avoid complications of unification, are presented as Fig. 1.

Table 1 presents unit root tests and Cochrane-based measures of persistence for the full sample period for each country. With the exception of the USA, the ADF unit-root test leads to a nonrejection of the null. Furthermore, a ranking of the persistence measures assigns the TS USA data with the lowest value. Such results are consistent with and mirror closely those of Cogley (1990).

Table 2 considers the effects of structural change on the ADF test statistic. As is well known, significant structural change will tend to bias the test statistic in favour of nonrejection of the unit-root null, see Perron (1989). With the exceptions of Italy,[2] the Netherlands, Norway and Sweden, a single significant discontinuity which overturns the unit-root null is identified for each country. Most of these discontinuities are associated with the World Wars although for

[2] For Italy, an 1896 trend change, 1918 jump and trend change, a 1939 crash and trend change and TSQ post 1939, gives an ADF(1) statistic of −6.136.

Table 1. *ADF and spectral-based measures of persistence (without discontinuities), log GDP per capita*

Country	Sample	lag	ADF	Persistence
Australia	1870–1993	0	−1.356	1.089
Austria	1870–1993	0	−1.663	0.840
Belgium	1870–1993	1	−1.247	1.706
Canada	1870–1993	1	−3.149	0.836
Denmark	1870–1993	0	−1.923	0.874
Finland	1870–1993	4	−2.103	1.471
France	1871–1993	1	−2.138	1.103
Germany	1871–1989	0	−2.087	0.656
Italy	1870–1993	1	−1.850	1.800
Japan	1885–1993	0	−1.246	1.621
Netherlands	1900–1993	1	−2.545	0.843
Norway	1870–1993	0	−1.624	1.541
Sweden	1870–1993	1	−1.823	2.188
UK	1870–1993	1	−2.104	0.951
USA	1870–1993	1	−4.132*	0.369

Notes: * Rejects the appropriate null at the 5% level based upon MacKinnon (1991). The degree of augmentation (lag) has been chosen primarily to remove residual serial correlation, but contingent on this, we include significant lagged dependent variables and choose on the basis of a Bayesian Information Criterion (BIC). The window size for the spectral density was 22 (except Germany, Japan and the Netherlands where it was 20) based upon the Chatfield rule, see Chatfield (1989).

Table 2. *Unit root tests, ADF (with discontinuities), log GDP per capita*

Country	1	Jump	Trend	j & t
Australia [year]	2	−3.420 (1891)	−4.447* (1931)	−4.629 (1925)
Austria	2	−3.401 (1913)	−4.632* (1945)	−5.130* (1944)
Belgium	2	−2.929 (1959)	−3.908 (1943)	−5.268* (1939)
Canada	2	−5.241* (1917)	−4.007 (1933)	−5.326* (1928)
Denmark	1	−4.122 (1958)	−3.930 (1941)	−5.552* (1939)
Finland	3	−3.781 (1958)	−4.127 (1924)	−5.247* (1913)
France	3	−3.877 (1959)	−4.400* (1942)	−5.500* (1939)
Germany	1	−5.203* (1946)	−4.066 (1930)	−4.980 (1946)
Italy	1	−3.278 (1918)	−3.703 (1944)	−3.989 (1939)
Japan	2	−3.303 (1959)	−3.454 (1945)	−5.104* (1944)
Netherlands	1	−3.746 (1930)	−3.745 (1943)	−4.465 (1944)
Norway	1	−3.043 (1944)	−3.383 (1940)	−3.531 (1916)
Sweden	4	−3.375 (1912)	−3.052 (1917)	−4.025 (1913)
UK	1	−3.505 (1918)	−4.677* (1931)	−6.084* (1918)

Notes. * Denotes significant at 5%. Figures in parentheses denote the year of the minimum t value. The terms 'jump', trend and j & t are defined in Section II.

Australia and Canada, see Oxley and Greasley (1995) country specific shocks are identified.[3]

Table 3 reports spectral-based measures of persistence for the full and subsample periods where the latter are determined on the basis of significant rejections of the null hypothesis of a unit root reported in Table 2. A number of tendencies can be observed.

For countries where the full sample DS results are overturned with the inclusion of discontinuities there is a general tendency for the second (and full sample) subperiod persistence measures to exceed the first. Exceptions to this tendency are Australia (marginal) Austria and Germany. There is also a tendency for second subperiod results, generally post-World War Two (WW2), to exceed full sample results. Exceptions here include Australia, where the break is identified in 1931, Austria, Finland (marginal) and Germany (marginal).

Considering all the subperiod results, and particularly those for the earlier period, the exceptionally low measure reported by Cogley for the USA seems less idiosyncratic. Belgium pre-1939; France pre-1939, and Japan pre-1944, exhibit lower persistence. The UK and Canada record measures close to 0.5.

Belgium, Italy, Japan, Norway and Sweden

Of the remaining results those for Belgium, Italy, Japan, Norway and Sweden have apparently high persistence (in excess of 1.5) and deserve closer inspection. It appears that a combination of (sometimes complex) breaks and nonlinearities often associated with periods of accelerating growth may partially explain these idiosyncratically high persistence results.

In the case of Italy a trend change in 1896, which coincides with a well documented industrial spurt (see Milward and Saul, 1977), a jump and trend change in 1918 and 1939 and a nonlinear term from 1940, leads to an ADF(1) statistic of −6.136 with all the intervention terms highly significant.

Table 4 presents measures of persistence for Italy for subperiods coinciding with significant discontinuities identified from the ADF procedure. The general tendency of earlier subperiod results to be considerably lower than latter period measures is repeated. The full sample measure of 1.8

[3] For Australia a two break model is favoured with breaks at 1891 and 1925. For Canada breaks at 1917 or 1928 lead to a rejection of the DS null.

Table 3. *Spectral-based measures of persistence (with and without discontinuities), log GDP per capita*

Country	Sample	Persist	k	t	Sample	Persist	k	t	Sample	Persist	k	t
Australia	1870–1993	1.089	22	124	1870–1931	0.919	16	62	1931–1993	0.869	16	63
Austria	1870–1993	0.840	22	124	1870–1945	0.960	18	76	1945–1993	0.644	14	49
Belgium	1870–1993	1.706	22	124	1870–1939	0.228	16	70	1939–1993	2.942	14	55
Canada	1870–1993	0.836	22	124	1870–1928	0.673	16	59	1928–1993	1.368	16	66
Denmark	1870–1993	0.874	22	124	1870–1939	0.412	16	70	1939–1993	0.944	14	55
Finland	1870–1993	1.471	22	124	1870–1913	1.018	14	44	1913–1993	1.465	18	50
France	1871–1993	1.103	22	123	1871–1939	0.238	16	69	1939–1993	1.572	14	55
Germany	1871–1989	0.656	20	119	1871–1946	1.282	18	76	1946–1989	0.648	14	44
Italy	1870–1993	1.800	22	124	–				–			
Japan	1885–1993	1.621	20	109	1870–1944	0.230	16	75	1944–1993	1.720	14	50
Netherlands	1900–1993	0.843	20	94	–				–			
Norway	1870–1993	1.541	22	124	–				–			
Sweden	1870–1993	2.188	22	124	–				–			
UK	1870–1993	0.951	22	124	1870–1918	0.550	14	49	1918–1993	1.162	18	76
USA	1870–1993	0.369	22	124	–				–			

Table 4. *Spectral-based measures of persistence, log Italian GDP per capita*

Sample	k	t	Persistence
1870–1993	22	124	1.800
1870–1896	10	27	0.296
1896–1918	10	23	0.772
1918–1939	10	22	1.888
1939–1993	14	55	1.502

coincides with the maximum identified for the 1918–39 subperiod with a tendency for measured persistence to decline post-1939. However, a significant post-1939 nonlinear time trend term casts doubt on the post-1939 and full sample measures which may be spuriously inflated.

In the case of Belgium it is not omitted nonlinear terms but a growth slowdown in 1973 which probably explains the high measure of persistence estimated at 2.942 for the 1939–93 period. For the period 1939–73 the spectral-based measure of persistence is 2.6 but for the period 1973–93 the measure declines to 0.698. Omitting the 1973 structural break leads to an overstating of post-WW2 persistence. A similar effect can be observed for Japan which also exhibits a 1973 slowdown in per capita output growth. For the period 1944–93 measured persistence is 1.720 but 1944–73 is the period of high persistence (1.781) whereas post-1973 this declines to 0.916.

Two of the highest full period measures of persistence are associated with Norway (1.541) and Sweden (2.188). Significant nonlinearities exist for the full sample in the case of Norway (coupled with a 1940 crash), see Fig. 2. For Sweden

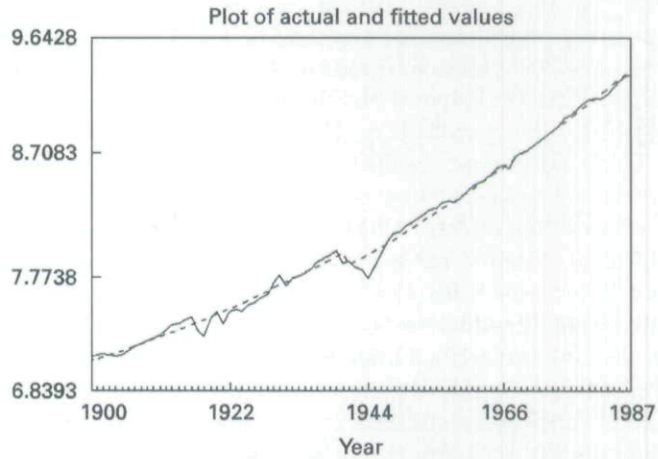

Fig. 2. *Plot of actual and fitted values:* (——) *LNDR,* (-----) *fitted.*

significant nonlinearities and a 1914 crash and 1973 trend break are significant in the ADF test procedure (see Fig. 3).[4] Such non-linearities undermine the use of Cochrane-type persistence measures, even over specific sub-samples because of the pervasiveness of the non-linear terms.

Subsample robustness: generating nonlinearities

The five specific countries identified above, Belgium, Italy, Japan, Norway and Sweden, highlight the importance of potential omitted nonlinearities and structural change. Furthermore, throughout the analysis we have assumed that discontinuous periods can be adequately joined at the breakpoint implying the validity of the full use of the available

[4] If include Time Squared (TSQ) and DU1940 in Norway, ADF(4) = −6.021. If include TSQ, DU1914, DT1973 in Sweden, ADF(4) = −6.680. If include DT1896, DT1918 DU1918, TSQ (post 1939) DT1939, DU1939, ADF(1) = −6.136, in Italy. If include DU1939, DT1939, DT1973, ADF(1) = −6.691, in Belgium.

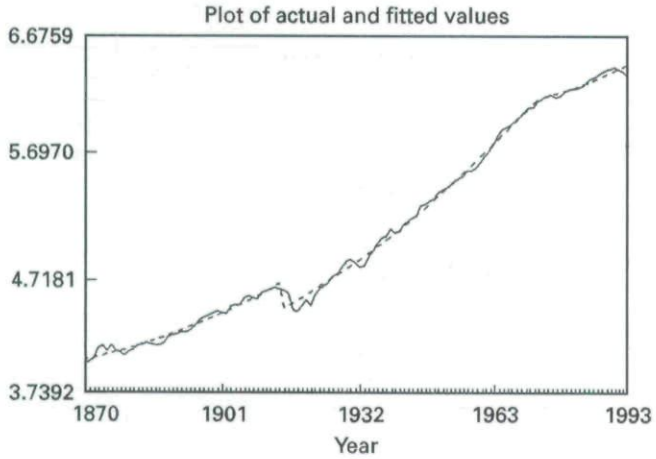

Fig. 3. *Plot of actual and fitted values:* (——) *LSWE*, (-----) *fitted.*

observations. De Long and Summers (1988) implicitly concede that two dis-joint sample periods exist for the USA when they report separate measures of 0.42, 1889–1929, and 0.82, 1947–87.[5] We will return to this particular point below. First the issue of the timing and sensitivity of the chosen sample periods is developed via Table 5.

Table 3 presented results for separate subperiod measure (where a breakpoint had been identified) on the basis of results obtained from ADF-type tests (supported by historical precedent). Table 5 considers the sensitivity of the second period results to the timing of the second period subsample. Identification of the timing of the initial break seems less controversial than the timing of the first observation of the second subperiod. In particular, should the second subperiod commence at the trough/peak of a major discontinuity or at a more representative ('normal') observation? Within the basis of the discussion so far, choosing an inappropriate ('nonrepresentative') start-point may manifest itself as a significant nonlinearity in the subsample. Via Table 5 and Table 3, we can observe that choice of second subperiod start points can be very important and explain some apparent outlying measures of persistence.

For Australia, Canada, and Denmark, the choice of second period start point has little effect on the measures reported in Tables 3 and 5. This might be expected as the nature of the breaks is generally relatively small. In the cases of Austria, Germany, and Japan, and to a lesser extent, Belgium, Finland, France and the UK, the timing of the second period start point can be of enormous importance. For Austria a 1945–93 measure of 0.644 emerges to be contrasted with measures in excess of 5.0 when the sample period is incremented with single observations to 1946–93, 1947–93, and 1948–93. Similar effects arise for Germany and Japan. The quantitative effects are somewhat smaller in other cases but again the breaks are often smaller. What seems to be happening here is that the choice of start point can significantly influence the nature of empirical nonlinearities in the data set.

This issue can be pursued further by including results for the USA which also exhibit similar sensitivity issues and which highlight the importance of the DeLong and Summers (1988) choice of subsamples.

The USA in more detail

So far we have excluded detailed discussion of the results for the USA. This stems in part from the TS (without breaks) ADF result. However, De Long and Summers (1988) suggest that post-WW2 persistence is higher in the USA than pre-war. Table 6 may help explain the reason. Table 6 presents subperiod results for USA persistence for all the periods identified in Table 3 for the 14 countries considered. This allows for common window size results for the specific countries to be compared with the USA subperiod result to eliminate any window size effect on the comparisons. It also provides an explanation of the De Long and Summers result. For subperiods which exclude WW2, persistence appears relatively low, as it does for the full sample where the war years enter near the middle of the sample. In contrast when the war years enter either at the beginning or the end of a sub-sample, measures rise by a factor of three.

It appears that when these untypically high levels of GDP per capita enter the sample at the beginning or end of the sample period they overwhelm the more typical observations in the measure of persistence. This coincides with the type of result reported in Table 5. Similarly the timing of the second subperiod (and the possibility of further breaks, i.e., a 1973 slowdown in GDP growth) affect the degree of measured persistence, see Table 6. This is particularly relevant when considering the De Long and Summers (1988) results based upon a choice of 1947–88. Table 6 shows that for most subperiods the full sample low persistence result is supported. Yet when WW2 years are included at the beginning or end of the sample the persistence measure rises considerably, see 1870–1944; 1870–1945; 1939–93; 1944–93; 1945–93, and 1946–93 with the spectral-based measure exceeding unity. When the sample period starts in 1947 or later the low persistence measure returns, see 1947–93, 1948–93. Furthermore, if the well-documented 1973 slowdown period is used as a breakpoint the end of WW2 effect exacerbates the 'excess' persistence measure, see 1945–73, 1946–73. Nevertheless the 1947–73 and 1948–73 results do show a tendency for persistence to rise above the pre-war levels, but not in excess of unity as would be indicated by the immediate WW2 start point. Post-1973, however, the measure declines to the Cogley average of approximately 0.5.

[5] Their measure is based upon GDP per worker not per capita.

Table 5. *Spectral-based measures of persistence, log GDP per capita*

Country	Sample	Persist	Sample	Persist	Sample	Persist
Australia	1932–1993	1.23	1933–1993	1.19	1934–1993	1.07
Austria	1946–1993	6.28	1947–1993	5.65	1948–1993	5.10
Belgium	1940–1993	2.63	1941–1993	2.02	1942–1993	1.73
Canada	1930–1993	1.41	1931–1993	1.19	1932–1993	1.19
Denmark	1940–1993	1.04	1941–1993	0.47	1942–1993	0.86
Finland	1914–1993	1.26	1915–1993	1.08	1916–1993	0.93
France	1940–1993	1.74	1941–1993	1.35	1942–1993	1.30
Germany	1947–1989	4.03	1948–1993	6.80	1949–1993	6.31
Japan	1945–1993	1.47	1946–1993	7.35	1947–1993	7.34
UK	1919–1993	1.12	1920–1993	0.63	1921–1993	0.58

Table 6. *Spectral-based measure of persistence, log GDP per capita, USA*

Sample	k	t	Persistence
1870–1913	14	44	0.338
1870–1918	14	49	0.284
1870–1928	16	59	0.223
1870–1931	16	62	0.471
1870–1939	16	70	0.461
1870–1944	18	75	1.403
1870–1945	18	76	1.142
1870–1946	18	77	0.488
1870–1947	18	78	0.432
1870–1948	18	79	0.434
1913–1993	18	81	0.569
1918–1993	18	76	0.612
1928–1993	16	66	0.815
1931–1993	16	63	0.747
1939–1993	14	55	1.212
1944–1993	14	50	1.168
1945–1993	14	49	1.614
1946–1993	14	48	1.310
1947–1993	14	47	0.425
1948–1993	14	46	0.481
1945–1973	10	29	1.973
1946–1973	10	28	1.645
1947–1973	10	27	0.778
1948–1973	10	26	0.877
1973–1993	10	21	0.546

Results for all countries: post 1950

Given the importance of choosing normal years as start points for any second period subsamples, and given the problems of comparing different country measures based upon different sample sizes and hence windows, Table 7 normalizes all countries' post-WW2 samples to commence at 1950, but permits a 1973 breakpoint. This allows some of the results derived above for Belgium, Italy, Japan, Norway and Sweden, where nonlinearities and further structural breaks were identified, to be extended to all countries under investigation.

Table 7. *Spectral-based measures of persistence, log GDP per capita*

Country	1950–93	1950–73	1973–93
Australia	2.34	2.25	1.04
Austria	4.82	2.04	1.12
Belgium	2.86	2.56	0.70
Canada	1.67	1.54	1.61
Denmark	1.47	1.14	0.53
Finland	2.68	0.68	2.89
France	7.43	1.08	2.81
Germany	6.22	5.27	0.60
Italy	5.64	2.08	1.27
Japan	7.77	1.32	0.92
Netherlands	2.92	0.57	1.29
Norway	1.37	0.58	1.95
Sweden	4.53	1.59	2.68
UK	0.77	0.42	0.73
USA	0.56	1.32	0.54

From Table 7 it can be seen with the exception of Canada the omission of a 1973 effect leads to dramatic effects on measures of persistence. Specifically the omission of a 1973 effect leads to, except the USA, Finland and Norway, a higher full sample, 1950–93 measure, than for the two subperiods, 1950–73, 1973–93. In many cases the relatively high measures of persistence are limited to the post-WW2, pre-1973, subsample. In all cases except the USA, significant 1973 trend changes and/or nonlinearities (identified by the significance of a post-WW2 TSQ term) were identified.

For Japan, Germany, Austria, Finland and Norway, the significance of TSQ appears particularly important in explaining the apparently high measures. For Canada, only the 1973 break appears significant, as it does in the case of Belgium. Japan also exhibits a well documented 1973 slowdown, exacerbating the nonlinear effects captured by TSQ. The omission of post-WW2 breaks and the apparent neglect of highly nonlinear elements in GDP per capita may lead to highly unstable and potentially erroneous measures of persistence.

IV. CONCLUSIONS

Since the work of Dickey and Fuller (1979) considerable attention has been directed towards testing for unit roots in economic time series. Cochrane (1988) highlighted the problems with some of the inferences which might be derived from such unit root tests and presented methods to measure the size of the random walk component. Generally, the Cochrane message has had little impact and few papers present measures of the degree of persistence. Exceptions include Cogley (1990), Leung (1992), Mayadunne *et al.* (1995) and Islam (1996). Leung (1992) considers the effects of structural breaks on such measures, albeit with the breakpoints arbitrarily chosen. DeLong and Summers (1988) compare two dis-joint sample periods, but do not consider the sensitivity of the periods chosen. Furthermore, Leung (1992) does not consider the effects of nonlinearities which may inflate Cochrane-type persistence measures, although he does warn of the dangers of ignoring discontinuities.

In two seminal contributions by DeLong and Summers (1988) and Cogley (1990) two apparently contradictory views on USA persistence have emerged. Noting the differences in the variable of interest, GDP per capita and per worker, it is interesting to note that Cogley (1990) chooses to ignore potential breaks in the series and furthermore, omits several 'difficult' countries from his results, i.e., Austria, Belgium, Finland, Japan, and Germany. DeLong and Summers (1988) however, consider dis-joint subsamples and in choosing 1947 as the initial post-WW2 observation and ignoring a possible 1973 effect, present a view of the USA's post-WW2 experience which might overstate persistence.

In this paper we have added to the debate via the careful consideration of a number of theoretical issues which may affect the Cochrane-type measures of persistence. In particular we have highlighted using an updated and extended version of the Cogley (1990) data set of GDP per capita, the need to consider the effects of structural change measured by either one-off changes in trend and/or intercept, and the significance of nonlinear elements. Furthermore, and closely linked to the issue of nonlinearities, we investigated the sensitivity of the results to the sample period chosen. We have demonstrated the apparent importance of the first and/or last few observations in the empirical sample when calculating measures of persistence.

Can any general conclusions be drawn? First, ignoring discontinuities and nonlinearities can affect the Cochrane-based measures of persistence and sometimes considerably. The general tendency for post WW2 persistence to increase relative to pre-WW2 periods appears evident in most of the series, although in many cases the relevance of the Cochrane-based measures is cast in doubt because of the highly nonlinear nature of the data set of the relevant subsample. Full sample estimates in these cases often underestimate the post-WW2 measures masking the tendency. Post-1973 measures appear to decline relative to those for the immediate post-WW2 period, suggesting a uniquely high period of persistence which has now typically been replaced by more 'normal' levels. For some data sets it appears that the average of subperiod results can lead to measures of persistence close to the full sample estimates. In many other cases this tendency does not emerge where the subperiod results are all below the full sample estimates. One feature that appears to be important is the choice of sample period. Using 'untypical' data at the beginning or end of a sample period appears to inflate the persistence measure – sometimes considerably. This appears in most cases to reflect either the creation of a subsequent (omitted) structural change or nonlinearity. Simply deleting certain untypical observations a la DeLong and Summers (1988) may be one option, alternatively careful modelling is required. Overall, the message when presenting Cochrane-based measures of persistence is a general one – if discontinuities exist in the data then check the robustness of the results. Nonlinearities, measured here by the significance of time-squared, need to be investigated as the Cochrane-type measure is based upon a linear expansion.

Considering the results for the USA in particular, for full and most subsamples including 1973–93, a low measure of persistence is observed. However, for the earlier subperiods this low measure is matched by several countries including Japan, Belgium and France, with the UK, Denmark and Canada close behind. For the USA subperiods which include the WW2 boom years, where the rate of growth exceeds the long-term average, much higher levels of persistence are observed. In part, the sensitivity of the persistence measure results to the sample period may explain the apparent contradictions in other authors' work.

In unit root tests ignoring a discontinuity can have serious consequences, see Perron (1989). The messages of Perron (1989) and Zivot and Andrews (1992) on the effects of discontinuities on Dickey–Fuller-type unit root tests are becoming more widely appreciated. The analogue for Cochrane-based measures is equally valid but much less well recognized.

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