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FURTHER EVIDENCE ON BREAKING TREND FUNCTIONS IN MACROECONOMIC VARIABLES

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RÉSUMÉ

Premièrement, cette étude réexamine les conclusions de Perron (1989) en ce qui a trait au fait que plusieurs des séries macroéconomiques sont mieux modélisées comme étant des fluctuations stationnaires autour d'une tendance déterministe si on admet la possibilité d'un changement d'ordonnée à l'origine en 1929 (un crash) et un changement dans la pente en 1973 (un ralentissement de la croissance). Contrairement à l'étude précédente, la date d'un changement possible n'est pas fixée *a priori*, mais est supposée inconnue. On considère ainsi différentes méthodes pour déterminer le point de discontinuité et les distributions échantillonnales asymptotiques et finies de la statistique correspondante. Une discussion détaillée à propos du choix du nombre de retards dans l'autorégression et son effet sur les valeurs critiques est aussi incluse. La plupart des rejets que l'on retrouvait dans Perron (1989) sont confirmés par cette approche. Deuxièmement, cet article examine des séries internationales trimestrielles sur le PNB (ou PIB) réel d'après-guerre pour les pays du G-7. D'autres séries sont analysées, dont la consommation réelle, desquelles, encore une fois, une forte évidence découle contre l'hypothèse de racine unitaire. Les résultats sont comparés à ceux de Banerjee, Lumsdaine et Stock (1992) et à ceux de Zivot et Andrews (1992). Par opposition aux résultats théoriques contenus dans ces articles, la distribution asymptotique des tests de type séquentiel est dérivée sans avoir recours à une exclusion des valeurs possibles des dates de changement près des bornes.

Mots-clés : test d'hypothèse, changement structurel, tendances stochastiques, tendances déterministes, expérience de simulation, racine unitaire.

ABSTRACT

Firstly, this study reexamines the findings of Perron (1989) regarding the claim that most macroeconomic time series are best construed as stationary fluctuations around a deterministic trend function if allowance is made for the possibility of a shift in the intercept of the trend function in 1929 (a crash) and a shift in slope in 1973 (a slowdown in growth). Unlike the previous study, the date of a possible change is not fixed *a priori* but is considered as unknown. We consider various methods to select the break points and the asymptotic and finite sample distributions of the corresponding statistics. A detailed discussion about the choice of the truncation lag parameter in the autoregression and its effect on the critical values is also included. Most of the rejections reported in Perron (1989) are confirmed using this approach. Secondly, this paper investigates an international data set of post-war quarterly real GNP (or GDP) series for the G-7 countries. A number of other series, including real consumption, are analyzed and strong evidence is again found against the unit root hypothesis. Our results are compared and contrasted to those of Banerjee, Lumsdaine and Stock (1992) and Zivot and Andrews (1992). In contrast to the theoretical results contained in these papers, we derive the limiting distribution of the sequential test without trimming.

Key words: hypothesis testing, structural change, stochastic trends, deterministic trends, simulation experiment, unit root.

1. INTRODUCTION.

In a previous paper, Perron (1989), we argued that many macroeconomic time series could be represented as stationary fluctuations around a deterministic trend function if allowance is made for a possible change in its intercept in 1929 (a crash) and in its slope in 1973 (a slowdown in growth). The test statistics were constructed by adding dummy variables for different intercepts and slopes, extending the standard Dickey–Fuller procedure. The asymptotic distribution theory underlying the critical values obtained under the different models assumed that the dating of the break points were known a priori, or more precisely, that the dates chosen were uncorrelated with the data.

This postulate has been criticized, most notably by Christiano (1992) who argued that the choice of these dates had to be viewed, to a large extent, as being correlated with the data. This is an important problem because both the finite sample and asymptotic distributions of the statistics depend upon the extent of the correlation between the choice of the break points and the data. There is a sense, as argued before, in which the choice of these dates can be regarded as independent of the data. First, the dates used in the previous study were chosen ex–ante and not modified ex–post. Secondly, these dates are related to exogenous events for which economic theory would suggest the effects that actually happened; e.g. the stock market crash of 1929 with the ensuing dismantle of the economic organization and the exogenous sudden change in oil prices with the resulting alteration of international economic coordination and policies.

In the sense described above the choice of the dates can be viewed as uncorrelated with the data. There is, however, a validity to the argument that it is only ex–post (after looking at the data) that we can say that the changes that followed these exogenous events actually occurred as predicted by the theory. Furthermore, many other exogenous events did not have the major impact that some theories would have predicted. In this sense, the choice of the break points must be viewed as being correlated, at least to some extent, with the data. To what extent is a difficult and practically impossible question to answer. At the very least the choices were not perfectly correlated with the data as no attempts were systematically made to maximize the chances that the unit root be rejected nor to find where, according to some test criteria, were the most likely dates of change.

While we still believe that the assumption about the exogeneity of the choice of the break points is a good first approximation to the true extent of the correlation with the data, it is useful to investigate how robust the results are to different postulates. The aim of this paper is to take the extreme view where the choice of the break points is effectively made to be perfectly correlated with the data. This case is instructive to study because if one can still reject the unit root hypothesis under such a scenario it must be the case that it would be rejected under a less stringent assumption.

We proceed as follows for the practical implementation. Again, as in the previous analysis, only one possible break point is allowed for any single series. This break point is first chosen such that the t–statistic for testing the null hypothesis of a unit root is smallest among all possible break points. Hence, using such a procedure, the choice of the break point is indeed perfectly correlated with the data. We also consider choosing the break point that corresponds to a minimal t–statistic on the parameter of the change in the trend function. This allows the mild a priori imposition of a one–sided change (i.e. a decrease in the intercept or the slope of the trend function). As will be seen, such a minor change allows substantial gains in power. We also investigate various issues regarding the choice of the truncation lag parameter in the estimated autoregressions and the effect on the critical values of using different criteria for choosing this lag length.

Our paper is closely related to and complements those of Banerjee, Lumsdaine and Stock (1992) and Zivot and Andrews (1992) in that similar procedures and series are analyzed. We extend their analysis in several directions. On a methodological level, we consider the asymptotic distribution of the sequential test based on the minimal value of the unit root tests over possible break points. We show the results of Zivot and Andrews (1992) to be valid without any trimming at the end points. The proof, which is of interest in itself, is based on projection arguments and introduces a method that can be applied to a variety of frameworks. Concerning the empirical results, our analysis is more extensive and shows that alternative procedures can lead to conclusions that are less favorable to the unit root than suggested in these two studies. We pay particular attention to the importance of the selection of the truncation lag on the outcome of the tests.

The paper is organized as follows. Section 2 reviews the statistical models and statistics involved. Section 3 discusses the asymptotic distribution of the test statistics under the null hypothesis of a unit root. Section 4 analyzes their finite sample distribution using simulation methods. Section 5 contains simulation experiments providing information about their and power under various data–generating processes. Section 6 presents the empirical results for the Nelson–Plosser (1982) data set. Section 7 analyzes an international data set of post–war quarterly real GNP series for the United States, Canada, Japan, the United Kingdom, West Germany, Italy and France. Section 8 presents additional evidence on other series such as consumption and on some series analyzed in Section 6 but obtained from alternative sources. Section 9 offers concluding comments. A mathematical appendix contains the derivation of the limiting distributions.

2. THE MODELS AND STATISTICS.

We briefly review, in this section, the models and statistical procedures that will be used to test for a unit root allowing for the presence of a change in the trend function occurring at most once. The reader is referred to Perron (1989) for more details. Throughout this paper, the time at which the change in the trend function occurs is denoted by T_b. The first model is concerned with the case where only a change in the intercept of the trend function is allowed under both the null and alternative hypotheses. Furthermore this change is assumed to occur gradually and in a way that depends on the correlation structure of the noise function. This was termed the "innovational outlier model" and can be succinctly represented, under the null hypothesis of a unit root, by:

$$y_t = y_{t-1} + b + \psi(L)(e_t + \delta D(T_b)_t), \tag{1.a}$$

where $D(T_b)_t = 1(t = T_b + 1)$ with $1(\cdot)$ the indicator function. The sequence $\{e_t\}$ is i.i.d. $(0, \sigma^2)$ and $\psi(L)$ is a possibly infinite lag polynomial in L (with $\psi(0) = 1$). Denoting by z_t the noise function of the series, we have $A(L)z_t = B(L)e_t$ and $\psi(L) = A(L)^{-1}B(L)$. It is assumed that the finite order polynomials $A(L)$ and $B(L)$ have all their roots outside the unit circle. The immediate impact of the change in the intercept is δ while the long run impact is $\psi(1)\delta$. Under the alternative hypothesis of stationary fluctuations, the model is:

$$y_t = a + ct + \Phi(L)(e_t + \theta DU_t), \tag{1.b}$$

where $DU_t = 1(t > T_b)$, $\Phi(L) = C(L)^{-1}D(L)$ with $C(L)$ and $D(L)$ finite order polynomials in L having all their roots outside the unit circle. The immediate impact of the change in the intercept of the trend function is θ while the long run impact is $\theta\Phi(1)$. Model (1.a) can be tested against model (1.b) using the t–statistic for testing $\alpha = 1$ in the following regression (with Δ the difference operator such that $\Delta y_t = y_t - y_{t-1}$):

$$y_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha y_{t-1} + \Sigma_{i=1}^{k} c_i \Delta y_{t-i} + e_t. \tag{1.c}$$

Regression (1.c), like the others that will follow, is in the spirit of the Dickey–Fuller (1979) and Said–Dickey (1984) methodology whereby autoregressive–moving average processes are approximated by autoregressive processes of order k.

Under the second model, both a change in the intercept and a change in the slope of the trend function are allowed at time T_b. With a similar notation, the model under the null and alternative hypotheses can respectively be represented as:

$$y_t = y_{t-1} + b + \psi(L)(e_t + \delta D(T_b)_t + \eta DU_t), \tag{2.a}$$

and

$$y_t = a + ct + \Phi(L)(e_t + \nu DU_t + \gamma DT_t), \tag{2.b}$$

where $DT_t = 1(t > T_b)t$. Model (2.a) can be tested against Model (2.b) using the t–statistic for the null hypothesis that $\alpha = 1$ in the following regression estimated by OLS:

$$y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_b)_t + \alpha y_{t-1} + \Sigma_{i=1}^{k} c_i \Delta y_{t-i} + e_t. \tag{2.c}$$

Under the third model, a change in the slope of the trend function is allowed but both segments of the trend function are joined at the time of break. Here the change is presumed to occur rapidly and corresponds to the "additive outlier model" in the terminology of Perron (1989). The model under the null and alternative hypotheses can be represented as:

$$y_t = y_{t-1} + b + \gamma DU_t + \psi(L)e_t, \tag{3.a}$$

and

$$y_t = a + ct + \gamma DT_t^* + \Phi(L)e_t, \tag{3.b}$$

where $DT_t^* = 1(t > T_b)(t - T_b)$. The null hypothesis given by Model (3.a) can be tested versus the alternative Model (3.b) using the following two–step procedure. First, the series is detrended using the following regression estimated by OLS:

$$y_t = \mu + \beta t + \gamma DT_t^* + \tilde{y}_t. \tag{3.c.i}$$

The test is then performed using the t–statistic for $\alpha = 1$ in the regression:

$$\tilde{y}_t = \alpha \tilde{y}_{t-1} + \Sigma_{i=1}^k c_i \Delta \tilde{y}_{t-i} + e_t. \qquad (3.c.ii)$$

We denote by $t_{\hat{\alpha}}(i,T_b,k)$ $(i = 1,2,3)$, the t–statistic for testing $\alpha = 1$ under model i with a break date T_b and truncation lag parameter k (i.e. the t–statistic for $\alpha = 1$ in regressions (1.c), (2.c) and (3.c.ii) for i = 1, 2, and 3, respectively). In the regressions described above, T_b and k treated as unknown. We next describe various data–dependent methods to select these values endogenously.

2.1 Methods to Choose the Break Date T_b.

We consider two methods to select T_b endogenously. As in Zivot and Andrews (1992) and Banerjee, Lumsdaine and Stock (1992), we first consider the procedure whereby T_b is selected as the value which minimizes the t–statistic for testing $\alpha = 1$ in the appropriate autoregression over some range for the break points. This procedure is in the tradition of tests for structural change when the date of the change is assumed unknown which often consider the maximum of a sequence of random variables as the statistic of interest. Such is the case, for instance, with the CUSUM test of Brown, Durbin and Evans (1975). The asymptotic distribution of $t^*_{\alpha}(1)$ and $t^*_{\alpha}(2)$ was studied by Zivot and Andrews (1992) under the condition that the range of possible values for the break point be restricted to some subset that excludes values at each end of the sample. In the next section, we show that the limiting distribution derived by Zivot and Andrews (1992) remains valid even without trimming. We define the statistics as $t^*_{\alpha}(i) = \mathrm{Min}_{T_b \in (k+1,\, T)}\, t_{\hat{\alpha}}(i,T_b,k)$ $(i = 1,2,3)$.

The following procedure is also analyzed. Instead of choosing the break point T_b so that $t_{\hat{\alpha}}(i,T_b,k)$ is minimized, it is chosen to minimize either $t_{\hat{\theta}}$, the t–statistic on the parameter associated with the change in the intercept (Model 1) or $t_{\hat{\gamma}}$, the t–statistic on the change in slope (Models 2 and 3). We denote the t–statistic on α (for a null hypothesis that $\alpha = 1$) obtained from such a procedure by $t^*_{\alpha,\theta}(1)$ for Model 1 and by $t^*_{\alpha,\gamma}(i)$ (i = 2, 3) for Models 2 and 3. More precisely, $t^*_{\alpha,\theta}(1) = t_{\hat{\alpha}}(1,T_b^*,k)$ where T_b^* is such that $t_{\hat{\theta}}(T_b^*) = \mathrm{Min}_{T_b \in (k+1,T)} t_{\hat{\theta}}(T_b,k)$, where again different specifications about the choice of k will be analyzed. The statistics $t^*_{\alpha,\gamma}(i)$ (i = 2, 3) are defined in an analogous fashion. This procedure is more akin to that used by Christiano (1992). The use of the t–statistic on the parameter associated with the change in the trend function allows the possibility of imposing the mild a priori restriction of a one-sided change. Hence by choosing T_b so that $t_{\hat{\theta}}$ in (1.c) or $t_{\hat{\gamma}}$ in (3.c.i) is minimized, we allow the date of the change in the trend function to be unknown but restrict the analysis to the cases of a "crash" or a slowdown in growth. We also discuss the case where the break point is selected using the same procedure without any a priori on the sign of the change. In this context the break date is selected using the maximum of the absolute value of $t_{\hat{\theta}}$ or $t_{\hat{\gamma}}$. The corresponding statistics are denoted by $t^*_{\alpha,|\theta|}(1)$ for Model 1 and $t^*_{\alpha,|\gamma|}(i)$ (i = 2, 3) for Models 2 and 3.

2.2 Methods to Select the Truncation Lag Parameter k.

There is now substantial evidence that using data-dependent methods to select the truncation lag parameter k leads to test statistics having better properties (stable size and higher power) than if a fixed k is chosen a priori (unless, of course, one happens to select that value of k which is best), see Ng and Perron (1994) and Perron and Vogelsang (1992). We consider, in this paper, two such data-dependent methods. The first is the one originally implemented by Perron (1989). It uses a general to specific recursive procedure based on the value of the t-statistic on the coefficient associated with the last lag in the estimated autoregression. More specifically, the procedure selects that value of k, say k^*, such that the coefficient on the last lag in an autoregression of order k^* is significant and that the coefficient in an autoregression of order greater than k^* is insignificant, up to some maximum order kmax selected a priori. In the simulations and empirical applications reported below, we use a two-sided 10% test based on the asymptotic normal distribution to assess the significance of the last lags. This procedure is denoted below a "t-sig".

Said and Dickey (1984) use yet a different method in their reported empirical application. It is based on testing whether additional lags are jointly significant using an "F-test" on the estimated coefficients. The exact procedure is as follows. First a maximum value of k, kmax, is specified. For a given value of T_b, the autoregression is estimated with kmax and (kmax – 1) lags. A 10% one-tailed F-test is used to assess whether the coefficient on the kmaxth lag is significant and if so, the value of k chosen is this maximum value. If not, the model is estimated with (kmax – 2) lags. The lag (kmax – 1) is deemed significant if either the F-test for (kmax – 2) versus (kmax – 1) lags or the F-test for (kmax – 2) versus kmax lags are significant based on the 10% critical values of the chi-square distribution. This is repeated by lowering k until a rejection that additional lags are insignificant occurs or some lower bound is attained. In the empirical applications, the lower bound is set to k = 1. This procedure is denoted below as "F-sig".

In the case where the noise function is assumed to be generated from a finite order autoregressive process, we can use results in Hall (1990) to show that the data–dependent methods described above lead to tests having the same asymptotic properties as would prevail if the true autoregressive order was selected to estimate the autoregression provided kmax is selected greater than the true value. In the more general case where moving–average components are permitted, Ng and Perron (1994) show that tests with such data dependent methods to select k have the same asymptotic distribution provided $kmax^3/T$ converges to 0.

We choose these "general to specific" procedures rather than methods based on information criteria, such as AIC, because the latter tend to select very parsimonious models leading to tests with sometimes serious size distortions and/or power losses. This finite sample performance is consistent with the finding of Ng and Perron (1994) who show that using an information criterion leads to a selected value of k that increases to infinity, as T increases, only at the very slow rate log(T). These theoretical results are in accord with various empirical results showing that using the AIC leads to very small values of k being selected (typically 0 or 1) and that oftentimes the estimated residuals exhibit serial correlation (see Perron (1994)).

3. THE ASYMPTOTIC DISTRIBUTION OF THE STATISTICS.

In this section, we consider the limiting distribution of the statistics. To simplify the derivations we suppose the data–generating process to be a random walk,

$$y_t = y_{t-1} + e_t, \qquad (t = 0, 1, ..., T) \tag{4}$$

where the errors e_t are martingale differences, and consider the statistics constructed with $k = 0$. Using arguments in Ng and Perron (1994), we can then state that the resulting limiting distribution remains the same when additional correlation is present and the statistics are constructed with one of the data–dependent method to select k. This holds provided $kmax^3/T \to 0$ as $T \to \infty$. This is the same strategy as used by Zivot and Andrews (1992) and Banerjee, Lumsdaine and Stock (1992). All statistics are asymptotically invariant to a change in intercept. Vogelsang and Perron (1994) show that they are not asymptotically invariant to a change in slope but that the asymptotic distribution corresponding to a zero change in slope is a better approximation to the finite sample distribution for values typically encountered in practice. The following Theorem concerning the asymptotic distribution of $t^*_{\alpha}(i)$ (i = 1, 2, 3) is proved in the appendix.

THEOREM 1: Let $\{y_t\}_0^T$ be generated by (4) and denote by " $\Rightarrow$ " weak convergence in distribution from the space D[0,1] to the space C[0,1] using the uniform metric on the space of functions on [0,1]. Then:

a) for i = 1, 2:

$$inf_{T_b \epsilon (1,T)} t_{\hat{\alpha}}(i, T_b, k=0) \Rightarrow inf_{\lambda \epsilon [0,1]} \int_0^1 W_i(r,\lambda) dW(r) / \left[\int_0^1 W_i(r,\lambda)^2 dr\right]^{1/2},$$

b) for Model 3: $inf_{T_b \epsilon (1,T)} t_{\hat{\alpha}}(i, T_b, k=0) \Rightarrow$

$$inf_{\lambda \epsilon [0,1]} \left[\int_0^1 W_3(r,\lambda) dW(r) - a \int_{\lambda}^1 (r-\lambda) W_0(r) dr \int_0^1 W_3(r,\lambda)^2 dr\right] / \left[\int_0^1 W_3(r,\lambda)^2 dr\right]^{1/2},$$

where $a = (\lambda^3(1-\lambda)^3/3)^{-1}$, $W_0(r)$ and $W_i(r,\lambda)$ are residuals from a projection of a standard Wiener process W(r) onto the subspace generated by the functions {1, r} (i = 0), {1, r, du(r,λ)} (i = 1), {1, r, du(r,λ), dt(r,λ)} (i = 2) and {1, r, dt*(r,λ)} (i = 3), with $du(r,\lambda) = 1(r > \lambda)$ and $dt^*(r,\lambda) = 1(r > \lambda)(r - \lambda)$.*

Theorem 1 differs from the results in Zivot and Andrews (1992) in two respects. First note that there is no need to introduce the hybrid metric considered in that paper. The weak convergence results hold under the uniform metric. This relaxation is achieved using arguments in Gregory and Hansen (1994) so that there is no need for a weak convergence result for DU_t or DT^*_t (appropriately normalized). The most important and novel aspect in which our result differs is that we do not require that the possible range of values for the break point be restricted to exclude the end points $\lambda = 0$ or 1. To achieve this relaxation, our proof is rather different and somewhat more involved than the one in Zivot and Andrews (1992) and is based on projection arguments. The intuition is quite simple. With a break at either end points, the regressions indeed exhibit perfect multicolinearity but the coefficient on the lagged dependent variable, α, is a linear combination of the parameter vector that is identifiable and estimable and its t–statistic is also well defined. In such cases, the regressions become equivalent to ones where no dummy is included and the standard limiting distribution of Dickey and Fuller (1979) applies.

This last result is important because it shows that it is unnecessary to use some arbitrary trimming near the end points, such as the 15% exclusion on both sides suggested by Banerjee, Lumsdaine and Stock (1992). The arguments presented in the proof of Theorem 1 can also be applied to other context such as the cointegration tests with regime shifts considered by Gregory and Hansen (1994).

We used simulation methods to obtain the percentage points of the asymptotic distributions described above. These were based on 10,000 replications using partial sums of i.i.d. N(0,1) random variables to approximate the Wiener process and 1,000 steps to compute the integrals. The critical values obtained are presented in the rows labelled "T = ∞" in panel A of Tables 1, 2 and 3.

This relaxation of the need for trimming at the end points does not appear to be possible for the tests whereby the break point is chosen with respect to the t–statistic on the coefficient of the intercept or slope change. The asymptotic distributions of $t^*_{\alpha,\theta}(1)$ and $t^*_{\alpha,|\theta|}(1)$ assuming the break point to be in some compact subset was derived in Banerjee, Lumsdaine and Stock (1992). The critical values are reproduced in Panels B and C of Table 1. Similar asymptotic results were obtained by Vogelsang and Perron (1994) for $t^*_{\alpha,\gamma}(i)$ and $t^*_{\alpha,|\gamma|}(i)$ $(i = 1, 2)$ and the critical values are in Panels B and C of Tables 2 and 3.

4. FINITE SAMPLE CRITICAL VALUES.

In this section we report simulation experiments to evaluate the finite sample distributions of the statistics under the null hypothesis of a unit root. Our aim is to assess the quality of the asymptotic approximation and to provide alternative sets of critical values when this approximation is inadequate. We consider the leading case of a random walk where the data are generated by:

$$y_t = y_{t-1} + e_t; \qquad\qquad y_0 = 0\,, \tag{5}$$

with $e_t \sim$ i.i.d. N(0,1). This setup allows us to assess the effects of different methods to select the truncation lag, especially those that are data–dependent. In the next section, we evaluate the size and power of the tests under various specifications for the value of the change in intercept and/or slope and the presence of additional correlation in the errors.

To assess the sensitivity of the distributions to the particular value of k used, we provide, for each sample size considered, simulated critical values for different specifications on the truncation lags, namely k = 0, 2, and 5 as well as chosen using the F–sig and t–sig methods. Given the nature of the data sets analyzed in later sections, we present critical values for the following sample sizes. For Model 1, T = 60, 80 and 100; for Model 2, T = 70 and 100; and for Model 3, T = 100, 150 and 200.

Each set of results was obtained using 2,000 replications of the t–statistic using the appropriate autoregression estimated from data generated by (5). The program was coded using the C language and N(0,1) random deviates were obtained from the routine RAN1 of Press et al. (1986). For the procedures where the choice of k is data dependent, kmax is set to 5 for purely computational reasons. Tables 1, 2 and 3 present the simulated critical values for Models 1, 2 and 3, respectively. Each table contains three panels: panel A for $t^*_\alpha(i)$ (i = 1, 2, 3), panel B for $t^*_{\alpha,\theta}(1)$ or $t^*_{\alpha,\gamma}(i)$ (i = 2, 3), and panel C for $t^*_{\alpha,|\theta|}(1)$ or $t^*_{\alpha,|\gamma|}(i)$ (i = 2, 3).

Table 1.A contains the critical values for $t^*_\alpha(1)$ where only a change in intercept is allowed. Upon comparison with the results in Perron (1989), it is readily seen that the critical values are much lower when T_b is allowed to be data dependent than when it is considered fixed. For example, with T = 100 and k = 0, the 5% critical value is –4.93 when minimizing over T_b as opposed to –3.76 when the date of the break is considered fixed at mid–sample. The critical values are fairly stable as k changes provided that k is held fixed when minimizing over T_b. In those cases where k is fixed, the asymptotic distribution is a good approximation to the finite sample distribution. The critical values for the test constructed with k chosen according to recursive F–tests on the coefficients of the lagged first differences are presented in the rows labelled k = k(F–sig). For example, the 5% point with T = 100 is –5.09. The critical values for the test constructed with k chosen according to a t–test on the last included lag in the autoregression are presented in the rows labelled k = k(t–sig). The resulting values are close to the values obtained using F–sig [1]. In those cases where a data dependent method is used to select k, the asymptotic approximation is not as good, indicating that the use of the asymptotic critical values would lead to tests that are liberal in finite samples.

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1 The simulated critical values involving a test of significance on the lagged first differences of the data are for tests of size 10%. We chose this value on the principle that it is safer to include extra lags to achieve the correct size in finite samples (at the expense of a loss in power). However, critical values with 5% tests were also computed and are not included since they are very similar to those with the 10% tests.

The critical values for $t^*_\alpha(2)$ (allowing both a change in slope and intercept) and $t^*_\alpha(3)$ (allowing only a change in slope) are presented in Tables 2.A and 3.A, respectively. The presentation of the results is as in Table I.A. The same general features hold when comparing different procedures and different sample sizes. While comparing the critical values for the three models, it is interesting to note that the highest critical values (in the left tail of the distribution) occur for Model 3. This is contrary to the fixed T_b case where the highest critical values correspond to Model 1.

Table 1.B presents the critical values of the statistic $t^*_{\alpha,\theta}(1)$ for Model 1 obtained with T_b chosen to minimize $t_{\hat{\theta}}$, the t–statistic on the parameter for the change in intercept. As can be seen from comparing the results in Table 1.A, the critical values obtained when choosing T_b this way are substantially smaller in absolute value. This is simply due to the a priori imposition of a one–sided change in the intercept of the trend function.

Tables 2.B and 3.B present the corresponding critical values for $t^*_{\alpha,\gamma}(2)$ and $t^*_{\alpha,\gamma}(3)$, obtained with T_b chosen to minimize $t_{\hat{\gamma}}$ in regressions (2.c) or (3.c.i). As stated earlier, this procedure does not impose any a priori restriction on the date of the change but restricts the change to be a decrease in slope (i.e. a one–sided structural change). Much of the same comments made with respect to the statistic $t^*_{\alpha,\theta}(1)$ apply to $t^*_{\alpha,\gamma}(2)$ and $t^*_{\alpha,\gamma}(3)$.

Panel C of Tables 1 to 3 consider the statistics based upon choosing the break date maximizing the absolute value of the t–statistic on the coefficient of the intercept or slope dummy. These statistics like $t^*_\alpha(i)$ ($i = 1, 2, 3$) do not impose any a priori condition on the sign of the change. Comparing the results with those in panels A, we see that for Models 1 and 3, the critical values in the left tail of the distribution are essentially the same between $t^*_\alpha(1)$ and $t^*_{\alpha,|\theta|}(1)$ and between $t^*_\alpha(3)$ and $t^*_{\alpha,|\gamma|}(3)$. Hence for Models 1 and 3, these two statistics are likely to have similar properties. Things are different for Model 2. The critical values in the left tail of the distribution are smaller (in absolute value) for $t^*_{\alpha,|\gamma|}(2)$ compared to $t^*_\alpha(2)$. Hence, one could expect the former to provide a more powerful test.

5. FINITE SAMPLE SIZE AND POWER SIMULATIONS.

This section presents finite sample size and power simulation results. The purpose is to determine the following, a) how size and power are affected by the choice of k in the presence of more general error processes, b) how size and power are affected by different values of the change in intercept and slope, and c) how power varies across procedures for choosing T_b. The focus of the simulations is placed on Models 1 and 3. The data generating process (DGP) used for Model 1 is of the form:

$$y_t = \theta DU_t + \delta D(T_b)_t + \alpha y_{t-1} + \Sigma_{i=1}^{4} \phi(i) \Delta y_{t-i} + (1 + \psi L) e_t, \tag{6}$$

where $e_t \sim$ i.i.d. N(0, 1) and $y_0 = e_0 = 0$. For Model 3, the DGP is of the form:

$$y_t = \gamma DT_t^* + \tilde{y}_t;\ \tilde{y}_t = \alpha \tilde{y}_{t-1} + \Sigma_{i=1}^{4} \phi(i) \Delta \tilde{y}_{t-i} + (1 + \psi L) e_t. \tag{7}$$

For the size simulations, $\alpha = 1$ and for power α is set to 0.8. The sample size for all simulations is T = 100 and 1,000 replications are used. Regressions were run for fixed k = 0, 1, ..., 5 and for k(F–sig) and k(t–sig) with kmax = 5. For fixed k, the 5% asymptotic critical values were used, and for k(F–sig) and k(t–sig), the appropriate 5% finite sample critical values for T = 100 were used. When the change in intercept or slope is non zero, the break date is $T_b = 50$ (at mid–sample). For Model 1, we used values of δ (under the null) and θ (under the alternative) of 0, 2, 5 and 10. For Model 3, we used values of γ at 0, .1, .3, .5 and 1. Seven different error specifications were used: 1) $\phi(i) = 0$ (i = 1, ..., 4) and $\psi = 0$; 2) $\phi(1) = .6$, $\phi(i) = 0$ (i = 2,3,4) and $\psi = 0$; 3) $\phi(1) = -.6$, $\phi(i) = 0$ (i = 2,3,4); $\phi(1) = .4$, $\phi(2) = .2$ and $\phi(3) = \phi(4) = \psi = 0$; 5) $\phi(1) = .3$, $\phi(2) = .3$, $\phi(3) = .24$, $\phi(4) = .14$ and $\psi = 0$; 6) $\phi(i) = 0$ (i = 1,2,3,4) and $\psi = .5$; 7) $\phi(i) = 0$ (i = 1,2,3,4) and $\psi = -.4$. Experiment (1) has i.i.d. errors. This specification is used to isolate the effects of choosing k too large. Experiment (2) has positive correlation in the errors and is quite common in empirical data. Experiment (3) has negative correlation in the errors. Experiments (4)–(5) have higher order correlation and are useful in isolating the effects of picking k too small. Finally, experiments (6) and (7) have MA(1) errors.

Due to space constraints, we only include results pertaining to $t_\alpha^*(3)$ in Table 4 (the full set of results is available on request). We begin by summarizing results pertaining to the choice of k. When k is chosen less than the true order of the process, substantial size distortions often occur. In most cases the exact size is much greater than the nominal size. If k is chosen at least as big as the true order of the process, the exact size is rarely greater than the nominal size. However, power is lost if the lag structure is over parameterized. When the k(t–sig) or k(F–sig) procedure is used to pick k, the exact size is close to the nominal size in all cases except when there is a negative MA component as in experiment (7). In this case the exact size is substantially inflated above the nominal size. Power using k(t–sig) or k(F–sig) is generally quite good. It is greater than when k is larger than the true order of the process and is nearly as high as when k is set to the true order in the case of autoregressive errors. Overall, the k(t–sig) and k(F–sig) procedures have good size and power properties and clearly dominate using a fixed k. The results indicate that tests based on the k(t–sig) procedure are slightly more powerful than those based on k(F–sig).

Consider now how a change in intercept or slope affects the exact size. The tests $t^*_\alpha(1)$ and $t^*_{\alpha,\theta}(1)$ become oversized as δ increases. For example consider experiment (1) for $t^*_\alpha(1)$ with k(t–sig); when $\delta = 0$ the size of the test is .047, when $\delta = 2$ it is .053, when $\delta = 5$ it is .096 and it rises to .486 when δ is as big as 10. The results in Tables 4 for $t^*_\alpha(3)$ concerning models with a change in slope γ show that changes in γ do not affect the size of the tests for the range of values considered. For $t^*_{\alpha,\gamma}(3)$, there are slight distortions in some cases as γ increases. Additional simulations revealed that larger values of γ induces substantial size distortions. The reader is referred to Vogelsang and Perron (1994) for a more detailed analysis on this issue. It is important to note here, though, that the magnitude of δ and γ where size distortions become a problem are of the order of 5 to 10 times the standard deviation of the errors for δ and at least 2 times the standard deviation of the errors for γ. For most macroeconomic time series (including those analyzed in later sections) intercept shifts are less than 5 standard deviations and slope changes are less than .7 standard deviations. Therefore distortions caused by large changes are not a problem in practice but care should be used if a series is suspected to have a very large intercept or slope change.

Consider now the effect on power of imposing the mild a priori condition on the sign of the change, i.e. comparing $t^*_\alpha(1)$ versus $t^*_{\alpha,\theta}(1)$ and $t^*_\alpha(3)$ versus $t^*_{\alpha,\gamma}(1)$. It is seen that power is generally higher when this condition is imposed.

As a final simulation experiment, we briefly analyze the test considered by Banerjee, Lumsdaine and Stock (1992) (BLS) and Zivot and Andrews (1992) (ZA) for the case of a joined segmented trend. They use the innovational outlier framework that does not allow for a change in slope under the null hypothesis. Namely, they use the regression:

$$y_t = \mu + \beta t + \gamma DT^*_t + \alpha y_{t-1} + \Sigma^k_{i=1} c_i \Delta y_{t-i} + e_t. \tag{8}$$

The statistic used is the minimal t–statistic on α for testing $\alpha = 1$. For illustration, we only consider the case where k is fixed at 4 as in BLS. We performed simulations using the same experiments as above with the DGP (7). The results are reported in Table 5. It is seen that this test, unlike the test $t^*_\alpha(3)$ based on the two step procedure, shows serious size distortions even with small non–zero values of the change in slope γ. For example, with i.i.d. errors, the size of this test is .320 (instead of the nominal 5%) when $\gamma = 1$, whereas the $t^*_\alpha(3)$ has an exact size close to 5% in this case. These results are of importance for the following reason. If one wants to treat a change in slope as a nuisance parameters (i.e. allow it under both the null and alternative hypotheses), the method suggested by BLS and ZA for the case of a joined segmented trend is inappropriate since a rejection could be due to either stationary fluctuations or a change in the slope of the trend function with a unit root process. The simulations reported here may also explain some simulation results in BLS. They document the fact that for a given change in slope the power of their test does not increase as α is further away from one. Given the results above, this may simply be due to the fact that most of the power for their parameter configurations is due to non–zero slope changes which have a first–order effect compared to gains in power caused by movements of α away from unity.

6. EMPIRICAL RESULTS FOR THE NELSON–PLOSSER DATA SET.

Table 6 presents the empirical results for the Nelson–Plosser (1982) series for which, in Perron (1989), Model 1 was the specification of interest. The series have an horizon which starts at different dates but which ends in 1970 for all series. The data set includes: Real GNP, Nominal GNP, Real per capita GNP, Industrial Production, Employment, GNP Deflator, Consumer Price Index, Nominal Wage, Money Stock, Velocity and Interest Rate. In Perron (1989) we claimed a rejection of the unit root for all these series except the Consumer Price Index, Velocity and Interest Rate. We also noted, however, that the unit root hypothesis could be rejected for the Consumer Price Index and Velocity series when considering the post–1929 sample using the usual Dickey–Fuller (1979) test statistic.

Table 6 presents results obtained by choosing the break point T_b in such a way that the t–statistic for testing that $\alpha = 1$ in regression (1.c) is minimized. For this set of results, kmax is specified to be 10 [2]. Results are presented for both cases where the truncation lag

[2] The choice of kmax is somewhat arbitrary. On the one hand, one would like a large value to have as unrestricted a procedure as possible. On the other hand, a large value of kmax yields problems of multicollinearity in the data and also a substantial loss of power. The choice of kmax was also set such that the estimated autoregressions did not show any sign of remaining correlation in the residuals as indicated by the Box–Pierce statistic. Most of the results are robust to alternative choices for kmax.

is selected using the F–sig or t–sig procedures. When both methods yield the same values only one regression is reported, otherwise separate regressions are listed. The statistics of most interest are the estimates of α and its t–statistic as well as three sets of p–values in the last three columns (reported to the nearest 1%). The first set of p–values is obtained using the asymptotic distribution. They are included because, as argued earlier, the asymptotic distributions of the statistics obtained under a data dependent method to select k are the same as in the fixed k case and may be more robust, for example to the presence of additional correlation, than the finite sample critical values. The second set of p–values corresponds to the critical values of the t–statistic when k and T_b are chosen according to the F–sig method. The last set of p–values are those corresponding to the t–sig method. The critical values used correspond to samples of size 60, 80 or 100 whichever is closest to the actual sample size.

The empirical results show that the unit root hypothesis can be rejected at the 5% significance level or better, under either scenario about the choice of k, for Real GNP, Nominal GNP [3], Industrial Production and Nominal Wages. For the Employment series, the finite sample p–value is .05 with the F–test and .09 with the t–test (the corresponding asymptotic p–values are .02 and .04, respectively). Hence the unit root is also rejected for the Employment series. The Real per capita GNP and Money Stock series present a more ambiguous case. When k is chosen with the F–sig procedure, the p–value for the Real per capita GNP series is .12 using the finite sample distribution and .06 using the asymptotic distribution. The corresponding figures are .14 and .08 for the Money Stock series. These values are marginal for a rejection at the 10% level. Nevertheless, to analyze whether these results are due to low power or are specific to the data series used here, Section 8 presents results that were obtained with similar series drawn from alternative sources.

The unit root hypothesis cannot be rejected for the Consumer Price Index, Velocity and Interest Rate series under any procedure. The choices of T_b and k obtained using the data dependent methods for choosing k are different but yield the same qualitative results.

The only series which offers a markedly different picture from the fixed T_b case is the

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3 For the Nominal GNP series, kmax was found binding in the sense that $k = 10$ was selected. Hence kmax was increased to 15 which again was found to be binding. We did not increase kmax further given the relatively few number of observations. Nevertheless, the conclusion is robust to basically any value of the truncation lag parameter k chosen.

GNP Deflator. With k chosen according to either significance criteria the p–value is .35 (.29 using the asymptotic distribution). Hence, for this series, it appears that the rejection of the unit root hypothesis reported in Perron (1989) is not robust to correlation between the choice of T_b and the data. It must be kept in mind, however, that the type of correlation assumed here is an extreme one and it may well be the case that the rejection would not hold under alternative specifications. Also of interest is the fact that the GNP Deflator series appears to behave in a manner similar to the Consumer Price Index: non–rejection of the unit root using the full sample but rejection using a standard Dickey–Fuller test on the post–1929 sample (see Table A.2 in Perron (1989)).

A comment is warranted about the choice of T_b selected according to these procedures. Except for the C.P.I, Velocity and Interest Rate series (for which the unit root is not rejected), the value of T_b is either 1929 or 1928. It is 1929 for the Nominal Wage and Money Stock series and 1928 for the others. While 1928 does not exactly correspond to the date specified in Perron (1989), the economic interpretation remains the same. The selection of 1928 is due to the presence of the dummy variable $D(T_b)_t$ in regression (1.c). Hence, 1928 is often chosen because the dummy variable takes value 1 in 1929 and offers some additional fit to the 1929 crash over what the change in the intercept can do alone.

Table 7 presents the results when T_b is chosen to minimize $t_{\hat{\theta}}$, the t–statistic on the parameter of the change in intercept, i.e. when imposing the one–sided restriction of a crash. We also considered the tests obtained without the imposition of the one–sided change, i.e. maximizing $|t_{\hat{\gamma}}|$. The results were qualitatively similar to those described above and are not reported. When a rejection of the unit root hypothesis occurred in Table 6, it does so again here and more strongly given that the tests have higher power. As was the case earlier, the unit root cannot be rejected for the GNP Deflator series. Hence, for this particular series, the earlier conclusion in Perron (1989) is not robust to allowing the date of the change to be unknown. The results in Table 7 offer, however, a different picture for three series. First, for the Employment series, the unit root hypothesis can be rejected at the 5% level (using any procedure) instead of 10% with the statistic $t^*_{\alpha}(1)$. More interestingly, the unit root hypothesis can now be rejected at the 10% level for the Real per capita GNP and Money Stock series. For example, the p–values under the F–sig procedure are .06 and .07, respectively.

We now turn to the analysis of the Common Stock Price and Real Wage series where Model 2 is specified, i.e allowing both a change in the intercept and the slope of the trend function. The procedures used and the presentation of the estimation results in Tables 8.A and 8.B follow our previous analysis of Model 1 (in Tables 6 and 7) except that kmax is now 5. Consider first the case where T_b is chosen to minimize the t–statistic on α. The date of break selected for the Common Stock Price series is 1928 (consistent with the imposition of 1929 as the break date in Perron (1989)). Both methods to choose the truncation lag yields the same model and test statistic with an asymptotic p–value of .02 and finite sample p–value of .04 for F–sig and .06 for t–sig. Similar results hold for the Real Wage series. The break date is 1939; the asymptotic p–value of the test is .03 and the finite sample ones are .07 with F–sig and .08 with t–sig.

Table 8.B presents results obtained when T_b is chosen maximizing $t_{\hat{\gamma}}$ or $|t_{\hat{\gamma}}|$, the t–statistic on the coefficient of the slope change. The results are quite interesting in that the unit root is strongly rejected using either method to select the truncation lag even without the a priori imposition on the sign of the change in slope. The selected break date is still 1939 for the Real Wage series but now 1936 for Common Stock Price.

To compare our results with those of Zivot and Andrews (1992), note first the methodological differences involved. First, we retained the one time dummy $D(T_b)_t$ in regressions (1.c) and (2.c); we consider the F–sig procedure to select the truncation lag as well as the t–sig procedure; we consider kmax = 5 instead of 10 for the Real Wages and Common Stock Price series; and we also consider the case where the break date is selected using a test of significance on the coefficient of the change in slope. For the series Real GNP, Nominal GNP, Industrial Production, Nominal Wages and Common Stock Prices our results concord with those of Zivot and Andrews (1992), namely a rejection of the unit root. Our results also show these rejections to be robust to alternative specifications for choosing the break date and the truncation lag (except for Nominal Wage using $t^*_\alpha(1)$ and F–sig). For the Employment series, our results allow a rejection at the 10% level using the finite sample critical values for the t–sig method when the break is selected minimizing the unit root statistic. This small decrease in p–value compared to Zivot and Andrews is basically due to the inclusion of the one–time dummy $D(T_b)_t$ in regression (1.c). However, our results show that a stronger rejection, at the 5% level, is possible using the F–sig method and that this rejection becomes even stronger if the mild a priori restriction of a one sided change is imposed (see Table 7). For the Real per capita GNP and Money Stock series, the results with $t^*_\alpha(1)$ and the t–sig procedure are similar to those in Zivot and Andrews (1992), namely p–values of .21 and .28. Using the F–sig procedure, the p–values are substantially reduced to .12 and .14, respectively. Imposing the sign of the change a priori allows a rejection at the 10% level for both series using F–sig and for Real per capita GNP using t–sig. The difference in our conclusion for the Real Wage series is due to the different choice of the upper bound kmax. Our results agree for non–rejections for the series GNP Deflator, C.P.I., Velocity and Interest Rate.

7. RESULTS WITH AN INTERNATIONAL DATA SET FOR POSTWAR REAL GNP.

This Section analyzes an international data set of postwar quarterly real GNP or GDP series. The countries analyzed and the type and sampling period of the series are the following: USA (GNP; 1947:1–1991:3); Canada (GDP; 1947:1–1989:1); Japan (GNP; 1957:1–1988:4); France (GDP; 1965:1–1988:3); Germany (GNP; 1960:1–1986:2); Italy (GDP; 1960:1– 1985:1); and the United Kingdom (GDP; 1957:1–1986:3). The data for USA are from the Citibase data bank and for Canada from the Cansim data bank. For Japan and France they are from the IFS data tape. The remaining series (U.K., Germany and Italy) are from Data Resources Inc. and are those used in Campbell and Mankiw (1989). All series are seasonally adjusted and at annual rates, except for the USA and the United Kingdom which are at quarterly rates. The plots of the logarithm of each series are presented in Figures 1 through 7. In these graphs the solid line is the estimated trend function allowing a one–time change in slope. The date of the change varies for each series and was selected using the $t_{\alpha}^{*}(3)$ test.

The results pertaining to the statistic $t_{\alpha}^{*}(3)$ are presented in Table 9. Using the asymptotic critical values (Table 3.A), the unit root is rejected, at close to the 5% level, for all series except Italy (for Canada this rejection is not robust when using the F–sig procedure). Using the finite sample critical values, the results are not, in general, as sharp. For Japan, the unit root is strongly rejected (p–values of .02 and .03). Indeed, the case of Japan is particularly striking in view of both the estimated t–statistic and the visual inspection of the graph of the series (Figure 3). The slope of the trend function has changed from 2.43% before 1971:3 to 1.01% after 1971:3 (a 58% decrease). More strikingly the actual series follows very closely this breaking trend function.

The results are not as clear for most of the other series but some interesting cases still emerge. For the United Kingdom, using the F–sig and t–sig procedures, the p–values are .07 and .08 respectively, allowing rejection of the unit root at the 10% level. The graph of the series is presented in Figure 7. The change in the slope of the trend function, with the break point evaluated at 1973:3, is still quite large showing a 54% reduction (with a quarterly growth rate at .74% before 1973:3 and .34% after).

The results for Canada, France, Germany and the United States are similar in terms of the t–statistics obtained. They range from –4.22 to –4.33 with finite sample p–values between .12 and .14 (this excludes the case of Canada with F–sig). While the unit root cannot be rejected at the 10% level, the results are not very much at odds with the hypothesis that the series can be construed as stationary fluctuations around a breaking trend function. Such is not the case, however, with the GDP series from Italy. Here the p–values are large enough to cast little doubts on the unit root.

It is interesting to look at the estimated change in the slope of the trend function and the dating of the break implied by the estimation procedure. The estimated percentage decrease in the rates of growth are as follows: USA, 37%; Canada, 36%; Japan, 58%; France, 66%; Germany, 56 %; Italy, 57%; U.K., 54%. These figures are indeed quite large and suggest, besides the unit root issue, that an important structural change has occurred. The dates of the break point are different for each country but are all close to the year 1973, associated with the first oil price shock. They vary between 1971:2 (USA) and 1976:3 (Canada). It is to be noted, however, that the statistical method used here is not directly geared at providing a consistent estimate of the date of change in the slope of the trend function. Hence, the break dates should be viewed as approximate.

As discussed in Section 5, using $t^*_{\alpha,\gamma}(3)$, which select T_b based on the minimal value of the parameter of the change in slope, is likely to allow tests with greater power. Results pertaining to this test are presented in Table 10. Indeed, it appears more powerful. Using the t–sig procedure, the p–values for the null hypothesis of a unit root are at most .11 for all countries except Italy. Using the F–sig procedure, the p–values are smaller than .10 for USA, Japan, France and the United Kingdom; they are .13 and .14 for Canada and Germany, respectively. These results show that a simple imposition of a one–sided downward change in slope (still with an unknown break point) is enough to warrant rejection of the unit root hypothesis at close to the 10% level for all countries except Italy.

We view these results, especially given the small span of the data, as substantial evidence against the unit root. It is indeed somewhat revealing to consistently obtain p–values in this range given the relatively low power of unit root tests when using a data series over a short span (see Perron (1991)). Given that the statistical procedure used is one where an extreme assumption is made about the correlation of the choice of the break point and the data (yielding a procedure with low power compared to the case where T_b is assumed fixed), we view these results as consistent with the hypothesis that the series are best characterized as stationary fluctuations around a breaking trend function with a change in slope near 1973.

To compare our results with those of Banerjee, Lumsdaine and Stock (1992) (BLS), we first note the main differences in the studies. First, the data used are slightly different in terms of both the sources and the horizon. Second, they use the one-step innovational outlier method which does not allow for a change in slope under the null hypothesis. Third, they use a fixed value for the truncation lag set at 4 for all countries and they note that the results are robust to setting this fixed value to 8 or to using an information criterion (AIC or Schwartz) to select the order. Using these different specifications they found little evidence against the unit root for all countries except Japan.

After several investigations using both types of methods applied to both data set [4], it turns out that the major factor responsible for the conflicting results is the method to choose the truncation lag. For example, our data dependent methods select k = 4 only for Japan for which we both reject the unit root. For the other countries the implied value of the selected truncation lag is different (in no case do our methods select k = 8 either). We believe our methods to select the truncation lag to be better for the purpose of the unit root tests for the following reasons documented in Ng and Perron (1994). First, fixing k to some arbitrary value can involve serious size distortions and/or power losses because the actual correlation structure of the data is not only unknown but is likely to be different across countries. As argued in Ng and Perron (1994), it is important to use a data dependent method for choosing the truncation lag when performing unit root tests. However, even among data dependent methods that implies asymptotically valid unit root tests, there are important differences between methods based on a general to specific approach and methods based on information criteria. In the context of a model where the noise component is an ARMA process, Ng and Perron (1994) show that the latter implies a sequences of selected values for k that increases with the sample size at a logarithmic rate, a very slow rate. The finite sample implication of this result is that methods based on information criteria will tend to select very low autoregressive orders. These implied parsimonious autoregressions will often not be enough to capture important serial correlation in the data and can lead to tests with size distortions and/or power losses.

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4 My thanks to Robin Lumsdaine for correspondance on this issue.

These theoretical issues are consistent with the empirical results of BLS who report values of k at 0 or 1 for all countries when using an information criterion. In no cases does our methods select such low values (except for Italy where we both agree for a non–rejection).

8. EMPIRICAL RESULTS FOR SOME ADDITIONAL SERIES.

Some additional series from alternative sources are analyzed in this Section. First, for the Real per capita GNP and Money Supply series, we use data sources other than the Nelson–Plosser data set. As discussed in Section 6, rejections of the unit root are borderline for these series when allowance is made for an unknown break point without imposing a one–sided change. To provide alternative evidence, we first present in Table 11 results related to the Friedman and Schwartz (1982) Real per capita GNP series for the same period (1909–1970), which is graphed in Figure 8. The results imply a maximum p–value of .03 under any method to select k and T_b [5], allowing an easy rejection of the unit root hypothesis for this series.

Consider now an alternative source for the money supply variable, the annual M2 series supplied in Balke and Gordon (1986) from 1869 to 1973, graphed in Figure 9. The results in Table 11 again show a strong rejection of the unit root with a p–value of at most .05 under any procedure.

Following the work of Hall (1978), much interest has been given to the time series behavior of consumption. To this effect, we analyze a data set consisting of historical series covering 1889 to 1973 for Nominal Consumption, Real Consumption, their per capita counterparts, the Consumption Price Index and also the Population series. These data are a subset of those used in Grossman and Shiller (1981). The graph of these series are presented in Figures 10 through 15. The results concerning the unit root tests are also presented in Table 11. For the Nominal and Real Consumption series the unit root can be rejected with a p–value less than .01 under any procedure. The series again exhibit a significant decline in their level in 1929. For the Nominal per capita Consumption series, a rejection is still possible with p–values at most .03 but the picture is different with the

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5 The rejection of the unit root for the Friedman and Schwartz series is robust to using the longer samples 1900–1973 and 1890–1973. It is not robust to using the whole sample 1869–1973. In the latter case, however, the unit root can be rejected using a standard Dickey–Fuller procedure without any allowance for a possible change in the trend function.

Real per capita Consumption series. Here the unit root cannot be rejected even though the break point is again associated with the 1929 crash. The results for the Consumption Price Index and Population series also imply a non–rejection of the unit root.

Table 12 presents the unit root tests obtained using the statistic $t^*_{\alpha,\delta}(1)$ which considers a one–sided structural change. The qualitative results are the same except for the Real per capita Consumption series where the p–values are now close to .10, casting some doubts on the unit root characterization for this series as well. The result for the Population series is interesting because it may explain why it is more difficult to reject the unit root for Real per capita GNP than for Real GNP itself. The result for the Consumption Price Index parallels the earlier results for the CPI and the GNP Deflator. It therefore appears that a broad class of price indices are characterized by a unit root.

9. CONCLUDING COMMENTS.

This paper documents the robustness of the results presented in Perron (1989). Unlike this previous study, we analyzed the case where the break date is explicitly correlated with the data and provided critical values to carry inference under a variety of procedures. This work is not intended as a substitute for the statistical procedures presented in that earlier paper but rather as a complement. Indeed, a case can often be made for using critical values that are based on the assumption of no correlation between the choice of the break point and the data. On the one hand, it may represent a close approximation to the actual extent of the correlation. On the other hand, each investigator may differ as to the amount of a priori information he or she is willing to incorporate into the analysis.

Another issue concerns the power of the tests. There appears to be a clear tradeoff between power and the amount of a priori information one is willing to incorporate with respect to the choice of the break point. The presumption is clearly that a procedure imposing no such a priori information, as the ones presented in this paper, has relatively low power. In this respect, the rejection of the unit root hypothesis, even when assuming a perfect correlation between the choice of the break point and the data, is quite strong.

APPENDIX: Proof of Theorem 1.

To simplify cross–references, we adopt the notation of Zivot and Andrews (1992), henceforth referred to as Z–A. Let $S_t = \Sigma_{j=1}^{t} e_j$ ($S_0 = 0$) and $X_T(r) = \sigma^{-1}T^{-1/2}S_{[Tr]}$, $(j-1)/T \leq r < j/T$ (for $j = 1, ..., T$), where $\sigma^2 = \lim_{T\to\infty} T^{-1}E(S_T^2)$ and $[\cdot]$ denotes the integer part of the argument. Since $\{e_t\}$ is i.i.d. with finite variance, we have $X_T(r) \Rightarrow W(r)$, where $\Rightarrow$ denotes weak convergence in distribution (from the space D[0,1] to the space C[0,1] using the uniform metric on the space of functions on [0,1]) with W(r) a standard Wiener process on [0,1]. Also, $\sigma_T^2 \equiv T^{-1}\Sigma_1^T e_t^2 \to_p \sigma^2$ where $\to_p$ denotes convergence in probability. Omitting the one–time dummy variable $D(T_b)_t$ (since it is asymptotically negligeable), we consider the following regressions:

$$y_t = \beta^i(\lambda)\, z_{tT}^i(\lambda) + \alpha^i(\lambda)\, y_{t-1} + e_t\,, \qquad (t = 1, ..., T), \tag{A.1}$$

for models i = 1, 2. The vector $z_{tT}^i(\lambda)$ encompasses the deterministic components of the model and depends explicitly on λ, the break fraction, and T, the sample size. For example, $z_{t,T}^1(\lambda)' = (1, t, DU_t(\lambda))$. Let $Z_T^i(\lambda,r) = \delta_T^i\, z_{[Tr],T}^i(\lambda)$ be a rescaled version with δ_T^i a diagonal matrix of weights. For example, $\delta_T^1 = \text{diag}(1, T^{-1}, 1)$. We also define the limiting functions $Z^1(\lambda,r) = (1, r, du(\lambda,r))'$ where $du(\lambda,r) = 1(r > \lambda)$, and $Z^2(\lambda,r) = (1, r, du(\lambda,r), dt^*(\lambda,r))'$ where $dt^*(\lambda,r) = 1(r > \lambda)(r - \lambda)$. Note that, as argued in Z–A, we do not have $Z_T^i(\lambda,r) \Rightarrow Z^i(\lambda,r)$ (i = A, C) as $T \to \infty$, using the uniform metric on the space of functions on D[0, 1]. The proof nevertheless remains valid without the need to introduce another metric to guarantee such convergence results. For simplicity, we henceforth drop the subscript denoting the model.

It is convenient to first transform (A.1) as follows. Let $Pz_T(\lambda) = [Pz_{1,T}(\lambda), ..., Pz_{T,T}(\lambda)]$ be the linear map projecting onto the space spanned by the columns of $z_T(\lambda)' = (z_{1,T}(\lambda), ..., z_{T,T}(\lambda))$. By definition $Pz_T(\lambda) = z_T(\lambda)(z_T'(\lambda)z_T(\lambda))^{-}z_T(\lambda)'$ where $(\cdot)^{-}$ denotes a g–inverse. Premultiplying by $Mz_T(\lambda) \equiv (I - Pz_T(\lambda))$, (A.1) can be written, in matrix notation, as:

$$Mz_T(\lambda)Y = \alpha(\lambda)Mz_T(\lambda)Y_{-1} + Mz_T(\lambda)e\,, \tag{A.2}$$

where $Y' = (y_1, ..., y_T)$, $Y'_{-1} = (y_0, ..., y_{T-1})$ and $e' = (e_1, ..., e_T)$. The t–statistic of interest can be written as

$$\inf_{\lambda\in[0,1]} t_{\hat{\alpha}}(\lambda) = [T^{-2}Y'_{-1}Mz_T(\lambda)Y_{-1}]^{-1/2}[T^{-1}Y'_{-1}Mz_T(\lambda)e]/s_T(\lambda) ,$$

where $s_T^2(\lambda) = T^{-1}(Y - \hat{\alpha}(\lambda)Y_{-1})'Mz_T(Y - \hat{\alpha}(\lambda)Y_{-1})$ with $\hat{\alpha}(\lambda)$ the OLS estimate of α in (A.2). We have:

$$T^{-2}Y'_{-1}Mz_T(\lambda)Y_{-1} = T^{-2}\Sigma_{t=1}^T\{y_{t-1} - z_{t,T}(\lambda)'[\Sigma_{s=1}^T z_{s,T}(\lambda)z_{s,T}(\lambda)']^-\Sigma_{s=1}^T z_{s,T}(\lambda)y_{s-1}\}^2$$

$$= T^{-1}\,\Sigma_{t=1}^T\{T^{-1/2}S_{t-1} - z_{t,T}(\lambda)'\,\delta_T[T^{-1}\Sigma_{s=1}^T\delta_T z_{s,T}(\lambda)z_{s,T}(\lambda)'\,\delta_T]^- T^{-1}\Sigma_{s=1}^T\delta_T z_{s,T}(\lambda)S_{s-1}\}^2 + o_{p\lambda}(1)$$

$$= \int_0^1\{\sigma X_T(r) - Z_T(\lambda,r)'[\int_0^1 Z_T(\lambda,s)Z_T(\lambda,s)'ds]^-\int_0^1 Z_T(\lambda,s)\sigma X_T(s)ds\}^2dr + o_{p\lambda}(1)$$

$$= \sigma^2\int_0^1\{X_T(r) - Pz_T(\lambda)X_T(r)\}^2dr + o_{p\lambda}(1) , \quad \text{(A.3)}$$

($o_{p\lambda}(1)$ denotes a random variable that converges in probability to 0 uniformly in λ) and:

$$Pz_T(\lambda)X_T(r) = Z_T(\lambda,r)'[\int_0^1 Z_T(\lambda,s)Z_T(\lambda,s)'ds]^-\int_0^1 Z_T(\lambda,s)X_T(s)ds.$$

Also,using developments as in Z–A,

$$T^{-1}Y'_{-1}Mz_T(\lambda)e$$
$$= T^{-1}\Sigma_{t=1}^T\{y_{t-1} - z_{t,T}(\lambda)'[\Sigma_{s=1}^T z_{s,T}(\lambda)z_{s,T}(\lambda)']^-\Sigma_{s=1}^T z_{s,T}(\lambda)y_{s-1}\}e_t + o_{p\lambda}(1)$$

$$= \sigma^2\int_0^1 X_T(r)dX_T(r) - \sigma^2\int_0^1 P_{Tz}(\lambda)X_T(r)dX_T(r) + o_{p\lambda}(1). \quad \text{(A.4)}$$

We can therefore express the t–statistic as a composite functional:

$\inf_{\lambda\epsilon[0,1]}t_{\hat{\alpha}}(\lambda) =$

$$g(X_T(r), \int_0^1 X_T(r)dX_T(r), Pz_T(\lambda)X_T(r), \int_0^1 Pz_T(\lambda)X_T(r)dX_T(r), s_T(\lambda)) + o_{p\lambda}(1),$$

where

$$g = h^*[h[H_1[X_T(r), Pz_T(\lambda)X_T(r)], H_2[\int_0^1 X_T(r)dX_T(r), \int_0^1 Pz_T(\lambda)X_T(r)dX_T(r)], s_T(\lambda)]],$$

with $h^*(m) = \inf_{\lambda\epsilon[0,1]}m(\lambda)$ for any real function $m = m(\cdot)$ on [0, 1]; and for any real functions $m_1(\cdot)$, $m_2(\cdot)$, $m_3(\cdot)$ on [0,1], $h[m_1(\lambda), m_2(\lambda), m_3(\lambda)] = m_1(\lambda)^{-1/2}m_2(\lambda)/m_3(\lambda)$. The functional H_1 and H_2 are defined by (A.3) and (A.4). The weak convergence results for each of the elements is contained in the following lemma.

Lemma A.1: The following convergence results hold jointly:

a) $X_T(r) \Rightarrow W(r)$;

b) $\int_0^1 X_T(r)dX_T(r) \Rightarrow \int_0^1 W(r)dW(r)$;

c) $Pz_T(\lambda)X_T(r) \Rightarrow Pz(\lambda)W(r) \equiv Z(\lambda,r)'[\int_0^1 Z(\lambda,s)Z(\lambda,s)'ds]^- \int_0^1 Z(\lambda,s)W(s)ds$;

d) $\int_0^1 P_T z(\lambda)X_T(r)dX_T(r) \Rightarrow \int_0^1 Pz(\lambda)W(r)dW(r)$;

e) $s_T^2(\lambda) = \sigma^2 + o_{p\lambda}(1)$.

Parts (a) and (b) are standard results, and part (e) follows using (c) and (d) and the fact that $T^{-1}\Sigma_1^T e_t \to_p \sigma^2$. To prove part (c), we start with the following Lemma which follows from Theorem 5.5 of Billingsley (1968).

Lemma A.2: $Pz_T(\lambda)X_T(r) \Rightarrow Pz(\lambda)W(r)$ *if* $X_T(r) \Rightarrow W(r)$ *and for any sequence of functions* $\{v_T(s)\}$ $(0 \leq s \leq 1)$ *approaching* $v(s)$, *we have:*

$$Pz_T(v_T(s)) \to Pz(v(s)), \tag{A.5}$$

where $Pz_T(v_T(s)) = Z_T(\lambda,r)'[\int_0^1 Z_T(\lambda,s)Z_T(\lambda,s)'ds]^- \int_0^1 Z_T(\lambda,s)v_T(s)ds$.

and $Pz(v(s)) = Z(\lambda,r)'[\int_0^1 Z(\lambda,s)Z(\lambda,s)'ds]^- \int_0^1 Z(\lambda,s)v(s)ds$.

We prove (A.5) in two steps. First, let

$$Pz(v_T(s)) = Z(\lambda,r)'[\int_0^1 Z(\lambda,s)Z(\lambda,s)'ds]^- \int_0^1 Z(\lambda,s)v_T(s)ds.$$

By the properties of projections in Hilbert spaces (e.g., Brockwell and Davis (1991, p. 52)):

$$Pz(v_T(s)) \to Pz(v(s)) \text{ if } v(s)_T \to v(s). \tag{A.6}$$

Now let

$$Pz_T(v(s)) = Z_T(\lambda,r)'[\int_0^1 Z_T(\lambda,s)Z_T(\lambda,s)'ds]^- \int_0^1 Z_T(\lambda,s)v(s)ds.$$

We need the following Lemma stated in Parthasarathy (1977, proposition 41.19).

Lemma A.3: Let $S_1 \subset S_2 \subset \ldots$ be an increasing sequence of subspaces in a Hilbert space $\mathcal{X}$ and let $S_\infty = \cup_i S_i$. Then $\lim_{T\to\infty} P(S_T)(x) = P(S_\infty)(x)$ for all x, where $P(S_T)(x)$ is the projection of x on the subspace S_T.

Lemma A.3 applied to our problem implies that

$$Pz_T(\lambda)(v(s)) \to Pz(\lambda)(v(s)), \tag{A.7}$$

since we can take $\mathcal{X} = D[0, 1]$ in which case $Z_T(\lambda,r) \in D[0, 1]$ and $Z(\lambda,r) \in C[0, 1] \subset D[0, 1]$.

Next we use the result that if for some sequence of random variables $\{X_T\}$ and $\{Y_T\}$ we have $X_T \Rightarrow X$ and $\| X_T - Y_T \| \to 0$ (under some P–measure), then $Y_T \Rightarrow X$ (under the same P–measure) (e.g., Billingsley (1968, Theorem 4.1) and Parthasarathy (1977, Corollary 51.3)). Let $X = Pz(v(s))$, $X_T = Pz_T(v(s))$ and $Y_T = Pz_T(v_T(s))$. Given (A.7), we only need to show that $\| Pz_T(v(s)) - Pz_T(v_T(s)) \| \to 0$. This follows easily since:

$$\begin{aligned} \| Pz_T(v(s)) - Pz_T(v_T(s)) \|^2 &= \| Pz_T(v(s) - v_T(s)) \|^2 \\ &= \| v(s) - v_T(s) \|^2 - \| v(s) - v_T(s) - Pz_T(v(s) - v_T(s)) \|^2 \\ &\le \| v(s) - v_T(s) \|^2 \to 0. \end{aligned}$$

This completes the proof of parts (c). To prove part (d), note that we have:

$$\int_0^1 P_T z(\lambda) X_T(r) dX_T(r) =$$

$$\int_0^1 Z_T(\lambda,r)' dX_T(r) \left[\int_0^1 Z_T(\lambda,s) Z_T(\lambda,s)' ds\right]^- \int_0^1 Z_T(\lambda,s) X_T(s) ds.$$

For concreteness consider model C where $Z_T(\lambda,s) \equiv (Z_{1,T}(s), Z_{2,T}(\lambda,s))'$ with $Z_{1,T}(s) \equiv (1, [Ts]/T)'$, $Z_{2,T}(\lambda,s) \equiv (1([Ts]/T > \lambda), 1([Ts]/T > \lambda)([Ts]/T - \lambda))'$ and $Z(\lambda,s) \equiv (Z_1(s), Z_2(s,\lambda))'$ with $Z_1(s) \equiv (1, s)'$, $Z_2(\lambda,s) \equiv (du(\lambda,s), dt^*(\lambda,s))'$. Also define $Z^*_{2,T}(s) = (1, [Ts]/T - \lambda)'$ and $Z^*_2(s) = (1, (s-\lambda))'$. Using arguments as in Gregory and Hansen (1994),

$$\int_0^1 Z_T(\lambda,s) Z_T(\lambda,s)' ds = \begin{bmatrix} \int_0^1 Z_{1,T}(s) Z_{1,T}(s)' ds & \int_0^1 Z_{1,T}(s) Z_{2,T}(\lambda,s)' ds \\ \int_0^1 Z_{2,T}(\lambda,s) Z_{1,T}(s)' ds & \int_0^1 Z_{2,T}(\lambda,s) Z_{2,T}(\lambda,s)' ds \end{bmatrix}$$

$$= \begin{bmatrix} \int_0^1 Z_{1,T}(s) Z_{1,T}(s)' ds & \int_\lambda^1 Z_{1,T}(s) Z^*_{2,T}(s)' ds \\ \int_\lambda^1 Z^*_{2,T}(s) Z_{1,T}(s)' ds & \int_\lambda^1 Z^*_{2,T}(s) Z^*_{2,T}(s)' ds \end{bmatrix} \Rightarrow \begin{bmatrix} \int_0^1 Z_1(s) Z_1(s)' ds & \int_\lambda^1 Z_1(s) Z^*_2(s)' ds \\ \int_\lambda^1 Z^*_2(s) Z_1(s)' ds & \int_\lambda^1 Z^*_2(s) Z^*_2(s)' ds \end{bmatrix}$$

$$= \begin{bmatrix} \int_0^1 Z_1(s) Z_1(s)' ds & \int_0^1 Z_1(s) Z_2(\lambda,s)' ds \\ \int_0^1 Z_2(\lambda,s) Z_1(s)' ds & \int_0^1 Z_2(\lambda,s) Z_2(\lambda,s)' ds \end{bmatrix} = \int_0^1 Z(\lambda,s) Z(\lambda,s)' ds.$$

Note that the result does require that $Z_T(\lambda,s) \Rightarrow Z(\lambda,s)$ under the uniform metric. Similarly, we have $\int_0^1 Z_T(\lambda,s) X_T(s) ds \Rightarrow \int_0^1 Z(\lambda,s) X(s) ds$. Finally, $\int_0^1 Z_T(\lambda,r)' dX_T(r) = T^{-1/2} \Sigma_1^T Z_T(\lambda,t/T) e_t = (T^{-1/2} \Sigma_1^T e_t,\ T^{-3/2} \Sigma_1^T t e_t,\ T^{-1/2} \Sigma_{T_b+1}^T e_t,\ T^{-3/2} \Sigma_{T_b+1}^T (t-T_b) e_t)$ $\Rightarrow (W(1),\ \int_0^1 r dW(r),\ W(1) - W(\lambda),\ \int_\lambda^1 (r-\lambda) dW(r)) = \int_0^1 P z(\lambda) W(r) dW(r)$. This completes the proof of part (d).

To complete the proof of the main result, we need to show continuity of the various functionals. Continuity of h^* and h is proved in Z–A.

Lemma A.2: The functions H_1 and H_2 defined by (A.3) and (A.4) are continuous at $(W(r), Pz(\lambda)W(r))$ and $(\int_0^1 W(r)dW(r), \int_0^1 Pz(\lambda)W(r)dW(r))$ with W–probability one.

Proof: Since H_1 and H_2 are continuous functions of their respective elements, the proof follows if each of the elements is bounded over [0,1] with W–probability one. $W(\cdot)$ is bounded with W–probability one and so is $\int_0^1 W(r)dW(r)$ as discussed in Z–A. Using arguments similar to those in Z–A, $\int_0^1 Pz(\lambda)W(r)dW(r)$ will be continuous if $Pz(\lambda)W(r)$ is continuous, i.e. if $\sup_{\lambda\in[0,1]}|Pz(\lambda)W(r)| < \infty$. We note that $Pz(\cdot)$ is a linear operator that maps an element on C[0, 1] (the Wiener process W(r) which is continuous) to a subspace defined by the functions $Z(\lambda,r)$. Continuity of $Pz(\lambda)W(r)$ follows since a linear projection map is bounded and continuous (see, e.g., Ash (1972), p. 130 and p. 148). □

It is useful to illustrate this result by way of an example. Consider Model A where $Z(\lambda, r) = (1, r, du(\lambda, r))$. Note that :

$$\int_0^1 Z(\lambda,s)Z(\lambda,s)'ds = \begin{bmatrix} 1 & 1/2 & (1-\lambda) \\ 1/2 & 1/3 & (1-\lambda^2)/2 \\ (1-\lambda) & (1-\lambda^2)/2 & (1-\lambda) \end{bmatrix}.$$

If $\lambda = 0$, $\int_0^1 Z(0,s)Z(0,s)'ds = \begin{bmatrix} 1 & 1/2 & 0 \\ 1/2 & 1/3 & 0 \\ 0 & 0 & 0 \end{bmatrix} \equiv A$,

and if $\lambda = 1$, $\int_0^1 Z(1,s)Z(1,s)'ds = \begin{bmatrix} 1 & 1/2 & 1 \\ 1/2 & 1/3 & 1/2 \\ 1 & 1/2 & 1 \end{bmatrix} \equiv B$.

A and B are obviously nonsingular, but a common g–inverse is given by

$$G = 12\begin{bmatrix} 1/3 & -1/2 & 0 \\ -1/2 & 1 & 0 \\ 0 & 0 & 0 \end{bmatrix}.$$

Since the choice of the g–inverse leaves a projection map unchanged, we have for $\lambda = 0, 1$:

$$Pz(\lambda)W(r) = Z^{\perp}(r)'[\int_0^1 Z^{\perp}(s)Z^{\perp}(s)'ds]^{-1}\int_0^1 Z^{\perp}(s)W(s)ds,$$

where $Z^{\perp}(r)' = (1, r)$, in which case the limiting distribution of $t_{\hat{\alpha}}(\lambda)$ $(\lambda = 0, 1)$ reduces to that in the case where no dummy for structural change is included.

The proof for Model 3 follows similar arguments and is therefore omitted. It uses the limiting distribution for fixed λ derived in Perron and Vogelsang (1993a,b) (see also Vogelsang (1993)).

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TABLE 1.A: Distribution of $t^*_{\hat{\alpha}}(1)$; Model 1; Choosing T_b minimizing $t_{\hat{\alpha}}$.

		1.0%	2.5%	5.0%	10.0%	50.0%	90.0%	95.0%	97.5%	99.0%
T = 60	k = 0	-5.61	-5.25	-4.93	-4.59	-3.62	-2.85	-2.64	-2.43	-2.22
	k = 2	-5.57	-5.18	-4.84	-4.54	-3.64	-2.82	-2.57	-2.31	-2.05
	k = 5	-5.49	-5.11	-4.88	-4.55	-3.61	-2.67	-2.41	-2.13	-1.77
	k = k(F-sig)	-5.83	-5.49	-5.21	-4.91	-3.91	-3.00	-2.70	-2.41	-1.96
	k = k(t-sig)	-5.92	-5.58	-5.23	-4.92	-3.91	-3.00	-2.74	-2.55	-2.25
T = 80	k = 0	-5.38	-5.11	-4.92	-4.62	-3.68	-2.85	-2.65	-2.43	-2.14
	k = 2	-5.47	-5.10	-4.85	-4.58	-3.64	-2.79	-2.57	-2.34	-2.05
	k = 5	-5.38	-5.08	-4.78	-4.53	-3.64	-2.76	-2.48	-2.27	-1.71
	k = k(F-sig)	-5.77	-5.35	-5.15	-4.84	-3.87	-2.96	-2.70	-2.41	-2.12
	k = k(t-sig)	-5.77	-5.31	-5.09	-4.84	-3.88	-2.95	-2.73	-2.55	-2.22
T = 100	k = 0	-5.49	-5.15	-4.93	-4.60	-3.70	-2.95	-2.67	-2.46	-2.26
	k = 2	-5.43	-5.12	-4.84	-4.56	-3.68	-2.87	-2.57	-2.33	-2.15
	k = 5	-5.40	-5.05	-4.85	-4.55	-3.68	-2.84	-2.56	-2.28	-1.95
	k = k(F-sig)	-5.70	-5.35	-5.09	-4.82	-3.89	-3.00	-2.74	-2.46	-2.22
	k = k(t-sig)	-5.70	-5.36	-5.10	-4.82	-3.87	-3.05	-2.75	-2.46	-2.22
T = ∞		-5.41	-5.02	-4.80	-4.58	-3.75	-2.99	-2.77	-2.56	-2.32

TABLE 1.B: Distribution of $t^*_{\alpha,\theta}(1)$; Model 1; Choosing T_b Minimizing $t_{\hat{\theta}}$

		1.0%	2.5%	5.0%	10.0%	50.0%	90.0%	95.0%	97.5%	99.0%
T = 60	k = 0	-5.37	-4.96	-4.55	-4.24	-3.19	-1.98	-1.54	-1.23	-0.64
	k = 2	-5.31	-4.88	-4.57	-4.26	-3.13	-1.85	-1.26	-0.74	0.03
	k = 5	-5.23	-4.87	-4.51	-4.20	-3.07	-1.43	-0.62	-0.05	0.74
	k = k(F-sig)	-5.58	-5.15	-4.88	-4.47	-3.33	-1.60	-0.84	-0.05	0.56
	k = k(t-sig)	-5.70	-5.21	-4.92	-4.53	-3.32	-1.79	-1.14	-0.35	0.42
T = 80	k = 0	-5.21	-4.91	-4.68	-4.34	-3.24	-2.11	-1.74	-1.28	-0.64
	k = 2	-5.28	-4.85	-4.60	-4.27	-3.20	-1.98	-1.48	-0.81	-0.05
	k = 5	-5.14	-4.84	-4.55	-4.21	-3.15	-1.73	-0.88	-0.26	0.69
	k = k(F-sig)	-5.50	-5.11	-4.85	-4.53	-3.33	-1.86	-1.06	-0.32	0.67
	k = k(t-sig)	-5.59	-5.09	-4.83	-4.54	-3.33	-1.92	-1.19	-0.46	0.34
T = 100	k = 0	-5.17	-4.90	-4.60	-4.30	-3.27	-2.16	-1.72	-1.33	-0.77
	k = 2	-5.21	-4.85	-4.58	-4.28	-3.21	-1.98	-1.49	-0.94	-0.33
	k = 5	-5.09	-4.85	-4.57	-4.27	-3.17	-1.81	-1.10	-0.58	0.15
	k = k(F-sig)	-5.42	-5.03	-4.80	-4.47	-3.33	-1.92	-1.33	-0.77	0.02
	k = k(t-sig)	-5.43	-5.05	-4.83	-4.50	-3.34	-2.02	-1.38	-0.84	-0.05
T = ∞		-5.15	-4.87	-4.64	-4.37	-3.39	-2.27	-1.85	-1.38	-0.70

TABLE 1.C: Distribution of $t^*_{\alpha,|\theta|}(1)$; Model 1; Choosing T_b maximizing $|t_{\hat{\theta}}|$.

		1.0%	2.5%	5.0%	10.0%	50.0%	90.0%	95.0%	97.5%	99.0%
T = 60	k = 0	-5.58	-5.17	-4.85	-4.50	-3.52	-2.46	-2.00	-1.59	-1.13
	k = 2	-5.51	-5.09	-4.78	-4.47	-3.51	-2.30	-1.70	-1.03	-0.37
	k = 5	-5.47	-5.09	-4.80	-4.48	-3.46	-1.78	-1.06	-0.26	0.41
	k = k(F-sig)	-5.77	-5.42	-5.13	-4.80	-3.70	-1.87	-1.19	-0.39	0.24
	k = k(t-sig)	-5.85	-5.51	-5.18	-4.83	-3.70	-2.14	-1.34	-0.55	0.05
T = 80	k = 0	-5.37	-5.08	-4.89	-4.57	-3.59	-2.49	-2.10	-1.76	-1.19
	k = 2	-5.46	-5.07	-4.83	-4.55	-3.54	-2.28	-1.74	-1.18	-0.45
	k = 5	-5.36	-5.05	-4.76	-4.48	-3.53	-2.07	-1.32	-0.71	0.26
	k = k(F-sig)	-5.75	-5.26	-5.06	-4.77	-3.71	-2.14	-1.42	-0.79	0.11
	k = k(t-sig)	-5.66	-5.29	-5.04	-4.78	-3.72	-2.28	-1.67	-0.96	-0.06
T = 100	k = 0	-5.47	-5.13	-4.89	-4.58	-3.62	-2.55	-2.18	-1.87	-1.21
	k = 2	-5.43	-5.07	-4.81	-4.52	-3.57	-2.41	-1.94	-1.41	-0.61
	k = 5	-5.40	-5.03	-4.80	-4.52	-3.56	-2.21	-1.58	-0.89	0.11
	k = k(F-sig)	-5.69	-5.34	-5.03	-4.75	-3.74	-2.33	-1.80	-1.20	-0.18
	k = k(t-sig)	-5.68	-5.36	-5.05	-4.77	-3.71	-2.40	-1.88	-1.21	-0.34
T = ∞		-5.34	-5.08	-4.84	-4.59	-3.74	-2.71	-2.35	-2.01	-1.54

TABLE 2.A: Distribution of $t^*_{\alpha}(2)$; Model 2; Choosing T_b minimizing $t_{\hat{\alpha}}$.

		1.0%	2.5%	5.0%	10.0%	50.0%	90.0%	95.0%	97.5%	99.0%
T = 70	k = 0	-6.65	-5.80	-5.43	-4.99	-3.98	-3.13	-2.89	-2.73	-2.52
	k = 2	-6.27	-5.67	-5.24	-4.87	-3.88	-3.06	-2.82	-2.63	-2.46
	k = 5	-6.21	-5.58	-5.27	-4.90	-3.89	-3.00	-2.74	-2.54	-2.34
	k = k(F-sig)	-6.22	-5.81	-5.52	-5.22	-4.21	-3.28	-3.00	-2.76	-2.54
	k = k(t-sig)	-6.32	-5.90	-5.59	-5.29	-4.24	-3.32	-3.08	-2.85	-2.67
T = 100	k = 0	-6.77	-5.78	-5.41	-5.02	-4.01	-3.22	-3.01	-2.76	-2.59
	k = 2	-6.30	-5.57	-5.29	-4.92	-3.95	-3.13	-2.89	-2.68	-2.48
	k = 5	-6.00	-5.53	-5.22	-4.89	-3.93	-3.07	-2.83	-2.66	-2.38
	k = k(F-sig)	-6.07	-5.72	-5.48	-5.17	-4.17	-3.29	-3.05	-2.83	-2.58
	k = k(t-sig)	-6.21	-5.86	-5.55	-5.25	-4.22	-3.35	-3.13	-2.85	-2.63
T = ∞		-5.57	-5.30	-5.08	-4.82	-3.98	-3.25	-3.06	-2.91	-2.72

TABLE 2.B: Distribution of $t^*_{\alpha,\gamma}(2)$; Model 2; Choosing T_b minimizing $t_{\hat{\gamma}}$.

		1.0%	2.5%	5.0%	10.0%	50.0%	90.0%	95.0%	97.5%	99.0%
T = 70	k = 0	-5.46	-4.89	-4.57	-4.21	-2.88	-1.57	-1.22	-0.90	-0.55
	k = 2	-5.46	-5.00	-4.57	-4.23	-2.80	-1.38	-0.99	-0.67	-0.16
	k = 5	-5.47	-4.95	-4.58	-4.19	-2.72	-1.28	-0.83	-0.43	0.04
	k = k(F–sig)	-5.77	-5.32	-4.95	-4.51	-2.92	-1.37	-0.93	-0.54	-0.02
	k = k(t–sig)	-5.77	-5.38	-4.98	-4.55	-3.04	-1.53	-1.10	-0.71	-0.27
T = 100	k = 0	-5.46	-4.96	-4.55	-4.22	-2.88	-1.62	-1.32	-0.95	-0.56
	k = 2	-5.27	-4.90	-4.59	-4.18	-2.82	-1.52	-1.09	-0.77	-0.42
	k = 5	-5.26	-4.90	-4.55	-4.14	-2.73	-1.42	-0.97	-0.63	-0.07
	k = k(F–sig)	-5.50	-5.16	-4.85	-4.47	-2.91	-1.50	-1.11	-0.73	-0.30
	k = k(t–sig)	-5.56	-5.23	-4.91	-4.47	-2.99	-1.55	-1.19	-0.78	-0.38
T = ∞		-5.28	-4.95	-4.62	-4.28	-2.94	-1.64	-1.33	-0.98	-0.59

TABLE 2.C: Distribution of $t^*_{\alpha,|\gamma|}(2)$; Model 2; Choosing T_b maximizing $|t_{\hat{\gamma}}|$.

		1.0%	2.5%	5.0%	10.0%	50.0%	90.0%	95.0%	97.5%	99.0%
T = 70	k = 0	-5.60	-5.29	-4.86	-4.55	-3.40	-2.23	-1.88	-1.60	-1.19
	k = 2	-5.69	-5.25	-4.93	-4.54	-3.35	-2.09	-1.75	-1.42	-0.91
	k = 5	-5.60	-5.24	-4.91	-4.54	-3.31	-2.00	-1.61	-1.26	-0.61
	k = k(F–sig)	-6.01	-5.56	-5.25	-4.88	-3.64	-2.17	-1.82	-1.37	-0.76
	k = k(t–sig)	-6.07	-5.61	-5.33	-4.94	-3.72	-2.28	-1.89	-1.50	-0.85
T = 100	k = 0	-5.59	-5.21	-4.93	-4.53	-3.42	-2.21	-1.88	-1.63	-1.29
	k = 2	-5.52	-5.21	-4.89	-4.54	-3.37	-2.07	-1.75	-1.48	-0.91
	k = 5	-5.45	-5.15	-4.88	-4.52	-3.30	-2.00	-1.66	-1.30	-0.82
	k = k(F–sig)	-5.72	-5.37	-5.14	-4.84	-3.54	-2.11	-1.76	-1.42	-0.89
	k = k(t–sig)	-5.86	-5.49	-5.19	-4.88	-3.60	-2.23	-1.87	-1.49	-0.95
T = ∞		-5.57	-5.20	-4.91	-4.59	-3.47	-2.15	-1.86	-1.59	-1.30

TABLE 3.A: Distribution of $t^*_{\alpha}(3)$; Model 3; Choosing T_b minimizing $t_{\hat{\alpha}}$.

		1.0%	2.5%	5.0%	10.0%	50.0%	90.0%	95.0%	97.5%	99.0%
T = 100	k = 0	-5.15	-4.81	-4.49	-4.19	-3.20	-2.37	-2.15	-1.99	-1.84
	k = 2	-5.04	-4.70	-4.43	-4.09	-3.12	-2.35	-2.17	-2.02	-1.89
	k = 5	-4.88	-4.57	-4.33	-4.05	-3.11	-2.39	-2.19	-2.09	-1.91
	k = k(F–sig)	-5.41	-4.99	-4.74	-4.44	-3.36	-2.53	-2.34	-2.21	-2.08
	k = k(t–sig)	-5.45	-5.11	-4.83	-4.48	-3.44	-2.60	-2.39	-2.22	-2.06
T = 150	k = 0	-4.99	-4.62	-4.40	-4.12	-3.12	-2.32	-2.15	-1.97	-1.84
	k = 2	-4.97	-4.66	-4.37	-4.06	-3.14	-2.35	-2.16	-2.01	-1.86
	k = 5	-4.85	-4.57	-4.38	-4.06	-3.11	-2.35	-2.19	-2.04	-1.91
	k = k(F–sig)	-5.19	-4.85	-4.59	-4.31	-3.32	-2.47	-2.29	-2.11	-1.96
	k = k(t–sig)	-5.28	-4.96	-4.65	-4.38	-3.33	-2.50	-2.30	-2.13	-1.93
T = 200	k = 0	-4.99	-4.62	-4.39	-4.12	-3.12	-2.32	-2.14	-1.96	-1.83
	k = 2	-4.97	-4.66	-4.37	-4.05	-3.13	-2.33	-2.14	-2.01	-1.83
	k = 5	-4.84	-4.57	-4.38	-4.06	-3.10	-2.34	-2.17	-2.03	-1.91
	k = k(F–sig)	-5.19	-4.84	-4.59	-4.30	-3.30	-2.46	-2.26	-2.09	-1.96
	k = k(t–sig)	-5.28	-4.96	-4.65	-4.38	-3.32	-2.48	-2.27	-2.10	-1.90
T = ∞		-4.91	-4.62	-4.36	-4.07	-3.13	-2.32	-2.12	-1.96	-1.78

TABLE 3.B: Distribution of $t^*_{\alpha,\gamma}(3)$; Model 3; Choosing T_b Minimizing $t_{\hat{\gamma}}$

		1.0%	2.5%	5.0%	10.0%	50.0%	90.0%	95.0%	97.5%	99.0%
T = 100	k = 0	-4.96	-4.56	-4.20	-3.87	-2.73	-1.68	-1.37	-1.09	-0.67
	k = 2	-4.75	-4.43	-4.08	-3.77	-2.68	-1.69	-1.35	-1.03	-0.69
	k = 5	-4.57	-4.31	-4.00	-3.65	-2.59	-1.70	-1.38	-1.14	-0.82
	k = k(F-sig)	-5.02	-4.69	-4.40	-3.99	-2.76	-1.76	-1.46	-1.12	-0.79
	k = k(t-sig)	-5.26	-4.82	-4.44	-4.07	-2.83	-1.76	-1.45	-1.12	-0.83
T = 150	k = 0	-4.71	-4.37	-4.10	-3.75	-2.69	-1.66	-1.31	-1.03	-0.52
	k = 2	-4.69	-4.35	-4.04	-3.76	-2.67	-1.66	-1.29	-0.94	-0.56
	k = 5	-4.56	-4.35	-4.02	-3.69	-2.62	-1.64	-1.32	-1.00	-0.64
	k = k(F-sig)	-4.89	-4.54	-4.27	-3.93	-2.74	-1.70	-1.33	-1.01	-0.64
	k = k(t-sig)	-5.00	-4.63	-4.36	-3.99	-2.78	-1.72	-1.40	-1.07	-0.49
T = 200	k = 0	-4.56	-4.32	-4.05	-3.71	-2.66	-1.51	-1.15	-0.79	-0.48
	k = 2	-4.56	-4.18	-3.98	-3.69	-2.63	-1.53	-1.14	-0.82	-0.52
	k = 5	-4.49	-4.23	-3.95	-3.62	-2.61	-1.53	-1.22	-0.86	-0.62
	k = k(F-sig)	-4.75	-4.43	-4.13	-3.79	-2.69	-1.53	-1.23	-0.90	-0.59
	k = k(t-sig)	-4.77	-4.50	-4.22	-3.83	-2.72	-1.57	-1.24	-0.96	-0.56
T = ∞		-4.67	-4.36	-4.08	-3.77	-2.65	-1.57	-1.22	-0.90	-0.49

TABLE 3.C: Distribution of $t^*_{\alpha,|\gamma|}(3)$; Model 3; Choosing T_b maximizing $|t_{\hat{\gamma}}|$.

		1.0%	2.5%	5.0%	10.0%	50.0%	90.0%	95.0%	97.5%	99.0%
T = 100	k = 0	-5.08	-4.75	-4.44	-4.15	-3.11	-2.18	-1.93	-1.68	-1.39
	k = 2	-5.02	-4.64	-4.34	-4.00	-3.01	-2.12	-1.85	-1.65	-1.32
	k = 5	-4.82	-4.48	-4.26	-3.90	-2.91	-2.08	-1.84	-1.61	-1.25
	k = k(F-sig)	-5.29	-4.87	-4.57	-4.27	-3.15	-2.19	-1.99	-1.69	-1.40
	k = k(t-sig)	-5.38	-5.02	-4.67	-4.36	-3.24	-2.28	-2.04	-1.75	-1.46
T = 150	k = 0	-4.98	-4.57	-4.37	-4.09	-3.06	-2.13	-1.83	-1.58	-1.21
	k = 2	-4.91	-4.62	-4.33	-3.99	-3.04	-2.11	-1.85	-1.55	-1.17
	k = 5	-4.80	-4.49	-4.29	-3.98	-3.00	-2.09	-1.83	-1.54	-1.13
	k = k(F-sig)	-5.15	-4.77	-4.49	-4.21	-3.15	-2.16	-1.89	-1.59	-1.19
	k = k(t-sig)	-5.23	-4.91	-4.57	-4.28	-3.18	-2.19	-1.92	-1.63	-1.30
T = 200	k = 0	-4.78	-4.53	-4.30	-3.98	-3.03	-2.10	-1.86	-1.62	-1.35
	k = 2	-4.85	-4.49	-4.18	-3.96	-2.99	-2.11	-1.84	-1.63	-1.25
	k = 5	-4.77	-4.44	-4.20	-3.91	-2.97	-2.11	-1.80	-1.58	-1.34
	k = k(F-sig)	-5.02	-4.75	-4.41	-4.10	-3.07	-2.11	-1.86	-1.63	-1.29
	k = k(t-sig)	-5.02	-4.75	-4.41	-4.17	-3.11	-2.15	-1.91	-1.68	-1.26
T = ∞		-4.87	-4.58	-4.34	-4.04	-3.08	-2.14	-1.87	-1.61	-1.30

TABLE 4: Finite Sample Size and Power Simulations; Model 3, $t_{\alpha}^{*}(3)$.

DGP: $y_t = \gamma DT_t^{*} + \tilde{y}_t$; $\tilde{y}_t = \alpha\tilde{y}_{t-1} + \Sigma_{i=1}^{4}\phi(i)\Delta\tilde{y}_{t-i} + (1 + \psi L)e_t$

$e_t \sim$ i.i.d. $N(0, 1)$; $T = 100$, $T_b = 50$; 2,000 replications; 5% nominal size; kmax = 5.

	Size ($\alpha = 1.0$)					Power ($\alpha = 0.8$)				
	γ					γ				
k	0.0	0.1	0.3	0.5	1.0	0.0	0.1	0.3	0.5	1.0
(1) $\phi(i) = 0.0$ (i=1,..,4), $\psi = 0.0$										
0	.049	.053	.055	.047	.036	.358	.365	.344	.331	.325
1	.044	.049	.048	.042	.037	.287	.299	.283	.277	.277
2	.045	.046	.048	.041	.040	.203	.215	.207	.199	.205
3	.038	.039	.042	.040	.047	.160	.177	.169	.165	.167
4	.035	.037	.039	.036	.041	.122	.129	.134	.130	.134
5	.035	.035	.039	.038	.039	.110	.123	.125	.116	.117
F–sig	.050	.054	.058	.055	.050	.235	.256	.244	.231	.233
t–sig	.049	.051	.058	.050	.045	.257	.278	.270	.259	.258
(2) $\phi(1) = 0.6$, $\psi = \phi(i) = 0.0$ (i=2,3,4)										
0	.000	.000	.000	.000	.001	.000	.000	.000	.000	.000
1	.058	.060	.056	.054	.062	.908	.903	.904	.902	.901
2	.046	.048	.049	.049	.055	.753	.758	.761	.753	.756
3	.045	.046	.040	.041	.045	.586	.592	.600	.594	.593
4	.037	.040	.034	.038	.044	.405	.426	.424	.417	.417
5	.033	.033	.034	.037	.045	.289	.302	.306	.305	.305
F–sig	.049	.051	.047	.047	.054	.676	.679	.679	.688	.693
t–sig	.049	.047	.038	.046	.049	.760	.773	.774	.778	.785
(3) $\phi(1) = -0.6$, $\psi = \phi(i) = 0.0$ (i=2,3,4)										
0	.858	.874	.873	.858	.848	.997	.997	.998	.998	.998
1	.051	.048	.043	.038	.034	.131	.132	.117	.114	.113
2	.046	.040	.040	.037	.034	.090	.100	.096	.094	.098
3	.044	.045	.041	.037	.039	.084	.098	.094	.099	.099
4	.034	.030	.033	.032	.033	.063	.074	.082	.083	.077
5	.033	.035	.038	.037	.039	.056	.073	.073	.075	.070
F–sig	.037	.040	.044	.042	.044	.091	.104	.104	.104	.100
t–sig	.039	.039	.042	.034	.037	.090	.105	.105	.104	.097
(4) $\phi(1) = 0.4$, $\phi(2) = 0.2$, $\psi = \phi(3) = \phi(4) = 0.0$										
0	.004	.004	.005	.004	.004	.001	.000	.000	.000	.000
1	.009	.008	.008	.007	.006	.432	.439	.428	.421	.424
2	.048	.051	.050	.049	.048	.756	.764	.765	.763	.760
3	.040	.042	.047	.044	.051	.598	.611	.611	.607	.602
4	.038	.039	.043	.044	.046	.413	.432	.438	.436	.421
5	.042	.043	.040	.040	.050	.300	.314	.318	.311	.313
F–sig	.040	.048	.049	.047	.050	.582	.593	.600	.591	.593
t–sig	.038	.040	.038	.040	.044	.607	.625	.626	.624	.620

TABLE 4 (cont'd): Finite Sample Size and Power Simulations; Model 3, $t^*_\alpha(3)$.

DGP: $y_t = \gamma DT^*_t + \tilde{y}_t$; $\tilde{y}_t = \alpha\tilde{y}_{t-1} + \Sigma_{i=1}^{4}\phi(i)\Delta\tilde{y}_{t-i} + (1 + \psi L)e_t$

e_t ~ i.i.d. N(0, 1) ; T = 100, T_b = 50; 2,000 replications; 5% nominal size; kmax = 5.

	Size ($\alpha = 1.0$)					Power ($\alpha = 0.8$)				
	γ					γ				
k	0.0	0.1	0.3	0.5	1.0	0.0	0.1	0.3	0.5	1.0
(5) $\phi(1) = .3, \phi(2) = .3, \phi(3) = .25, \phi(4) = .14, \psi = 0.0$										
0	.108	.107	.106	.105	.100	.000	.000	.000	.000	.000
1	.001	.001	.001	.001	.000	.033	.036	.033	.033	.034
2	.002	.002	.001	.001	.000	.566	.568	.569	.571	.578
3	.033	.033	.031	.031	.034	.877	.881	.878	.876	.873
4	.071	.073	.078	.077	.080	.904	.898	.900	.897	.899
5	.051	.051	.054	.051	.051	.762	.774	.771	.770	.769
F–sig	.048	.046	.050	.048	.051	.857	.863	.858	.858	.856
t–sig	.038	.037	.039	.037	.037	.855	.859	.864	.860	.859
(6) $\psi = 0.5, \phi(i) = 0.0$ (i=1,..,4)										
0	.002	.003	.003	.003	.001	.008	.010	.007	.007	.005
1	.150	.142	.160	.160	.156	.548	.552	.566	.545	.542
2	.021	.021	.028	.028	.020	.096	.100	.110	.103	.103
3	.052	.050	.057	.061	.059	.190	.202	.215	.203	.202
4	.034	.035	.033	.033	.034	.101	.105	.117	.111	.113
5	.038	.037	.039	.035	.041	.102	.111	.128	.119	.117
F–sig	.053	.055	.062	.065	.061	.206	.213	.222	.210	.209
t–sig	.067	.066	.070	.069	.064	.244	.258	.271	.257	.262
(7) $\psi = -0.4, \phi(i) = 0.0$ (i=1,..,4)										
0	.734	.738	.739	.717	.691	.996	.996	.996	.995	.995
1	.202	.210	.206	.189	.174	.751	.763	.727	.720	.715
2	.079	.087	.075	.074	.074	.395	.406	.383	.385	.376
3	.044	.048	.050	.046	.046	.252	.263	.254	.251	.245
4	.036	.037	.034	.033	.036	.159	.173	.163	.164	.156
5	.031	.031	.033	.032	.035	.131	.144	.139	.138	.135
F–sig	.145	.148	.156	.148	.149	.456	.476	.452	.451	.442
t–sig	.235	.247	.257	.251	.248	.631	.658	.646	.643	.647

TABLE 5: Finite Sample Size and Power Simulations; Model 3, BLS method.
DGP: $y_t = \gamma DT_t^* + \tilde{y}_t$; $\tilde{y}_t = \alpha\tilde{y}_{t-1} + \Sigma_{i=1}^{4}\phi(i)\Delta\tilde{y}_{t-i} + (1 + \psi L)e_t$
$e_t \sim$ i.i.d. N(0, 1) ; T = 100, T_b = 50 ; 2,000 replications; 5% nominal size; k = 4.

	Size ($\alpha = 1.0$)					Power ($\alpha = 0.8$)				
	γ					γ				
	0.0	0.1	0.3	0.5	1.0	0.0	0.1	0.3	0.5	1.0
(1) $\phi(i) = 0.0$ (i=1,..,4), $\psi = 0.0$										
	.073	.071	.089	.131	.320	.149	.173	.205	.292	.706
(2) $\phi(1) = 0.6$, $\psi = \phi(i) = 0.0$ (i=2,3,4)										
	.091	.090	.080	.091	.126	.452	.468	.479	.494	.650
(3) $\phi(1) = -0.6$, $\psi = \phi(i) = 0.0$ (i=2,3,4)										
	.067	.068	.093	.190	.660	.089	.124	.174	.347	.899
(4) $\phi(1) = 0.4$, $\phi(2) = 0.2$, $\psi = \phi(3) = \phi(4) = 0.0$										
	.090	.089	.082	.094	.130	.445	.470	.484	.501	.650
(5) $\phi(1) = .3$, $\phi(2) = .3$, $\phi(3) = .25$, $\phi(4) = .14$, $\psi = 0.0$										
	.168	.173	.166	.169	.172	.910	.917	.919	.923	.922
(6) $\psi = 0.5$, $\phi(i) = 0.0$ (i=1,..,4)										
	.068	.067	.074	.087	.165	.140	.146	.178	.188	.421
(7) $\psi = -0.4$, $\phi(i) = 0.0$ (i=1,..,4)										
	.072	.075	.111	.208	.690	.208	.239	.309	.551	.974

TABLE 6: Empirical Results, Nelson – Plosser Data; $t^*_{\hat{\alpha}}(1)$, kmax = 10.

Regression: $y_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha y_{t-1} + \Sigma_{i=1}^{k} c_i \Delta y_{t-i} + e_t$.

Series	Sample	T	T_b	k	$t_{\hat{\theta}}$	$\hat{\alpha}$	$t_{\hat{\alpha}}$	p–value (asy)	p–value (F–sig)	p–value (t–sig)
Real GNP	1909–1970	62	1928	9	–5.13	.190	–5.93	<.01	<.01	
			1928	8	–4.79	.267	–5.50	<.01		.03
Nominal GNP [a]	1909–1970	62	1928	11	–6.34	.404	–8.16	<.01		<.01
			1928	15	–5.94	.497	–6.21	<.01	<.01	
Real per Capita GNP	1909–1970	62	1928	9	–3.73	.313	–4.81	.06	.12	
			1928	7	–3.31	.484	–4.51	.13		.21
Industrial Production	1860–1970	111	1928	8	–5.18	.272	–6.01	<.01	<.01	<.01
Employment	1890–1970	81	1928	8	–3.42	.586	–5.14	.02	.05	
			1928	7	–3.11	.650	–4.91	.04		.09
GNP Deflator	1889–1970	82	1928	5	–3.28	.783	–4.14	.29	.35	.35
C.P.I.	1860–1970	111	1939	5	2.00	.948	–3.09	.88	.88	.88
Wages	1900–1970	71	1929	7	–4.32	.619	–5.41	<.01		.02
			1929	9	–4.10	.635	–4.62	.10	.16	
Money Stock	1889–1970	82	1929	7	–2.80	.783	–4.69	.08	.14	
			1927	6	–2.50	.831	–4.30	.21		.28
Velocity	1869–1970	102	1949	8	2.95	.830	–2.81	.95	.94	
			1946	0	3.24	.858	–3.29	.81		.81
Interest Rate	1900–1970	71	1965	3	3.86	.934	–1.35	>.99	>.99	
			1963	3	3.44	.928	–1.35	>.99		>.99

[a]: For Nominal GNP, kmax = 15 (See footnote 3).

Table 7: Empirical Results; Nelson–Plosser Data Set; Model 1.

$t^*_{\hat{\alpha},\hat{\theta}}(1)$; Choosing T_b minimizing $t_{\hat{\theta}}$; kmax = 10.

Regression: $y_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha y_{t-1} + \Sigma_{i=1}^{k} c_i \Delta y_{t-i} + e_t$.

Series	T_b	k	$t_{\hat{\theta}}$	$\hat{\alpha}$	$t_{\hat{\alpha}}$	p–value (asy)	p–value (F–sig)	p–value (t–sig)
Real GNP	1928	9	–5.13	.190	–5.93	<.01	<.01	
	1928	8	–4.79	.267	–5.50	<.01		.02
Nominal GNP[a]	1929	11	–6.78	.231	–7.86	<.01		<.01
	1928	15	–5.94	.497	–6.21	<.01	<.01	
Real Per Capita GNP	1928	9	–3.73	.313	–4.81	.03	.06	
	1928	7	–3.31	.484	–4.51	.07		.10
Industrial Production	1928	8	–5.18	.272	–6.01	<.01	<.01	<.01
Employment	1928	8	–3.42	.586	–5.14	.01	.02	
	1928	7	–3.11	.650	–4.91	.02		.04
GNP Deflator	1919	5	–3.51	.886	–3.24	.58		.54
	1919	9	–3.61	.829	–3.87	.27	.28	
C.P.I.	1919	5	–3.12	.982	–1.16	.98	.96	.96
Wages	1929	7	–4.32	.619	–5.41	<.01		.01
	1929	9	–4.10	.635	–4.62	.05	.08	
Money Stock	1929	7	–2.80	.783	–4.69	.04	.07	
	1928	6	–2.63	.824	–4.28	.12		.15
Velocity	1880	5	–2.74	.928	–1.62	.96	.93	
	1880	0	–2.46	.897	–2.43	.87		.83
Interest Rate	1920	0	–4.16	1.058	1.16	>.99		>.99
	1918	0	–3.59	1.079	2.08	>.99	>.99	

[a]: For Nominal GNP, kmax = 15 (see footnote 3).

TABLE 8.A: Empirical Results, Nelson – Plosser Data; $t^*_{\hat\alpha}(2)$; kmax = 5 .

Regression: $y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_b)_t + \alpha y_{t-1} + \Sigma_{i=1}^{k} c_i \Delta y_{t-i} + e_t$.

Series	Sample	T	T_b	k	$\hat\theta$	$\hat\beta$	$\hat\gamma$	$\hat\alpha$	$t_{\hat\alpha}$	p-value (asy)	p-value (F-sig)	p-value (t-sig)
Common Stock Prices	1871–1970	100	1928	1	–1.36 (–4.92)	.0065 (4.43)	.0141 (4.61)	.716	–5.50	.02	.04	.06
Real Wages	1900–1970	71	1939	3	–.25 (–2.65)	.0086 (5.26)	.0047 (3.38)	.390	–5.41	.03	.07	.08

TABLE 8.B: Empirical Results, Nelson – Plosser Data; Model 2.

$t^*_{\alpha,\gamma}(2)$ and $t^*_{\alpha,|\gamma|}(2)$; Choosing T_b maximizing $t_{\hat\gamma}$ or $|t_{\hat\gamma}|$; kmax = 5 .

Regression: $y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_b)_t + \alpha y_{t-1} + \Sigma_{i=1}^{k} c_i \Delta y_{t-i} + e_t$.

								Two-sided, max $\lvert t_{\hat\gamma}\rvert$			One sided; max $t_{\hat\gamma}$		
Series	T_b	k	$\hat\theta$	$\hat\beta$	$\hat\gamma$	$\hat\alpha$	$t_{\hat\alpha}$	p-value (asy)	p-value (F-sig)	p-value (t-sig)	p-value (asy)	p-value (F-sig)	p-value (t-sig)
Common Stock Prices	1936	3	–2.10 (–4.86)	.0094 (4.93)	.0268 (4.91)	.553	–5.49	.01	.02	.02	<.01	.01	.01
Real Wages	1939	3	–.25 (–2.65)	.0086 (5.26)	.0047 (3.38)	.390	–5.41	.02	.04	.04	<.01	.02	.02

TABLE 9: Empirical Results, $t^*_{\alpha}(3)$; Model 3, International Data Set.

Regression: $y_t = \mu + \beta t + \gamma DT^*_t + \tilde{y}_t$; $\tilde{y}_t = \alpha\tilde{y}_{t-1} + \Sigma^k_{i=1} c_i \Delta\tilde{y}_{t-i} + e_t$.

Series	Sample	T	T_b	k	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\alpha}$	$t_{\hat{\alpha}}$	p-value (asy)	p-value (F-sig)	p-value (t-sig)
USA ; GNP (kmax = 5)	47:1–91:3	179	71:2	2	.0088 (95.7)	–.0024 (–13.03)	.900	–4.22	.07	.12	.14
Canada ; GDP (kmax = 5)	47:1–89:1	169	76:3	3	.0122 (205.3)	–.0044 (–21.18)	.818	–4.26	.06		.13
			76:3	4	.0122 (205.3)	–.0044 (–21.18)	.830	–3.78	.19	.26	
Japan ; GNP (kmax = 5)	57:1–88:4	128	71:3	4	.0243 (211.9)	–.0142 (–76.11)	.650	–5.14	<.01	.02	
			71:4	4	.0241 (212.0)	–.0141 (–75.19)	.659	–5.11	<.01		.03
France[a]; GDP (kmax = 10)	65:1–88:3	95	74:2	3	.0127 (106.7)	–.0080 (–46.93)	.679	–4.33	.05	.12	.13
Germany ; GNP (kmax = 10)	60:1–86:2	106	73:4	7	.0104 (74.0)	–.0058 (–21.92)	.624	–4.33	.05	.12	.13
Italy ; GDP (kmax = 10)	60:1–85:1	101	72:4	3	.0125 (67.5)	–.0071 (–20.91)	.755	–3.90	.15	.25	
			73:2	1	.0123 (70.9)	–.0071 (–21.37)	.814	–3.31	.40		.56
U.K. ; GDP (kmax = 10)	57:1–86:3	118	73:3	7	.0074 (73.5)	–.0040 (–18.96)	.552	–4.62	.02	.07	.08

[a] A dummy taking value 1 in 1968:2 was included for France to take account of the general strike in May 1968.

TABLE 10: Empirical Results; International Data Set; Choosing T_b minimizing $t_{\hat{\gamma}}$.

Regression: $y_t = \mu + \beta t + \gamma DT_t^* + \tilde{y}_t \,;\, \tilde{y}_t = \alpha \tilde{y}_{t-1} + \Sigma_{i=1}^{k} c_i \Delta \tilde{y}_{t-i} + e_t$.

				With t–test on last lag				With F–test on additional lags			
	T_b	$\hat{\beta}$	$\hat{\gamma}$	k	$t_{\hat{\alpha}}$	p–value (asy)	p–value (t–sig)	k	$t_{\hat{\alpha}}$	p–value (asy)	p–value (F–sig)
U.S.A. (kmax = 5)	69:2	.0089 (88.9)	–.0024 (–13.34)	2	–4.16 (.900)	.04	.06	2	–4.16 (.900)	.04	.05
Canada (kmax = 5)	77:1	.0121 (209.2)	–.0045 (–21.24)	3	–4.25 (.817)	.03	.06	4	–3.77 (.829)	.10	.13
Japan (kmax = 5)	71:3	.0243 (211.9)	–.0142 (–76.11)	3	–4.99 (.685)	<.01	.02	4	–5.14 (.650)	<.01	<.01
France (kmax = 10)	74:1	.0129 (106.7)	–.0081 (–47.76)	2	–4.26 (.705)	.03	.07	3	–4.16 (.685)	.04	.07
Germany (kmax = 10)	73:1	.0106 (72.7)	–.0058 (–22.55)	4	–4.00 (.705)	.06	.11	6	–3.82 (.681)	.09	.14
Italy (kmax = 10)	73:3	.0122 (72.3)	–.0072 (–21.51)	1	–3.26 (.817)	.25	.33	1	–3.26 (.817)	.25	.29
United Kingdom (kmax = 10)	73:2	.0074 (72.6)	–.0040 (–18.98)	7	–4.61 (.554)	.01	.04	7	–4.61 (.554)	.01	.03

Note: t–statistics are in parentheses except for the columns labelled $t_{\hat{\alpha}}$ where the entries are the estimates of α.

TABLE 11: Empirical Results, Additional Series; $t^*_\alpha(1)$, kmax = 12.

Regression: $y_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha y_{t-1} + \Sigma_{i=1}^{k} c_i \Delta y_{t-i} + e_t$.

Series	Sample	T	T_b	k	$t_{\hat{\theta}}$	$\hat{\alpha}$	$t_{\hat{\alpha}}$	p–value (asy)	p–value (F–sig)	p–value (t–sig)
Real per Capita GNP (FS)	1909–1970	62	1928	11	–4.74	.202	–5.42	<.01	.03	.03
M2	1869–1973	105	1929	12	–4.21	.720	–4.69	.08	.13	.14
Nominal Consumption	1889–1973	85	1928	11	–6.00	.579	–6.78	<.01		<.01
			1928	12	–5.65	.614	–5.70	<.01	.01	
Real Consumption	1889–1973	85	1928	11	–5.96	.202	–6.45	<.01	<.01	
			1929	11	–5.78	.109	–6.19	<.01		<.01
Nominal Per Capita Cons.	1889–1973	85	1928	12	–5.12	.613	–5.25	.02	.03	.03
Real Per Capita Consumption	1889–1973	85	1928	12	–4.14	.174	–4.49	.13	.20	
			1928	10	–3.69	.371	–4.54	.12		.19
Consumption Price Index	1889–1973	85	1929	8	–3.77	.709	–4.71	.07	.13	
			1919	10	–3.86	.810	–4.34	.19		.26
Population	1889–1973	85	1917	11	3.35	.933	–4.82	.05	.10	
			1923	10	–1.94	.948	–3.48	.70		.71

Table 12: Empirical Results; Additional Series; Model 1.

$t^*_{\alpha,\theta}(1)$; Choosing T_b minimizing $t_{\hat{\theta}}$; kmax = 12.

Regression: $y_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha y_{t-1} + \Sigma_{i=1}^k c_i \Delta y_{t-i} + e_t$.

Series	T_b	k	$t_{\hat{\theta}}$	$\hat{\alpha}$	$t_{\hat{\alpha}}$	p–value (asy)	p–value (F–sig)	p–value (t–sig)
Real Per Capita GNP (F–S)	1928	11	–4.74	.202	–5.42	<.01	.01	.02
M2	1929	12	–4.21	.720	–4.69	.04	.07	.07
Nominal Consumption	1928	11	–6.00	.579	–6.78	<.01		<.01
	1928	12	–5.65	.614	–5.70	<.01	<.01	
Real Consumption	1928	11	–5.96	.202	–6.45	<.01	<.01	
	1929	11	–5.78	.109	–6.19	<.01		<.01
Nominal Per Capita Consumption	1928	12	–5.12	.613	–5.25	<.01	.02	.02
Real Per Capita Consumption	1928	12	–4.14	.174	–4.49	.07	.11	
	1928	10	–3.69	.371	–4.54	.06		.10
Consumption Price Index	1919	10	–3.86	.810	–4.34	.11	.14	.14
Population	1925	8	–2.36	.966	–2.39	.88	.81	.82

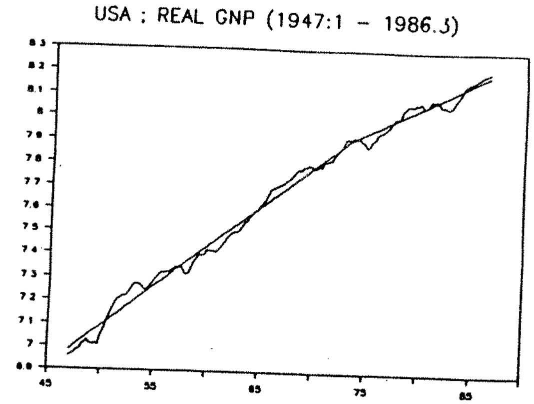

Figure 1

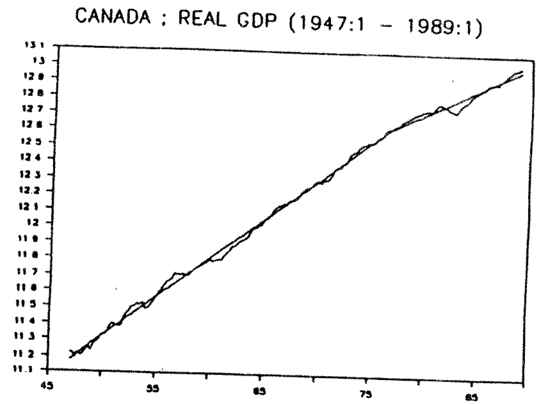

Figure 2

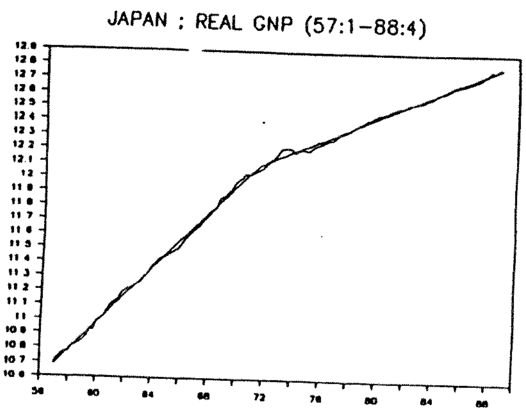

Figure 3

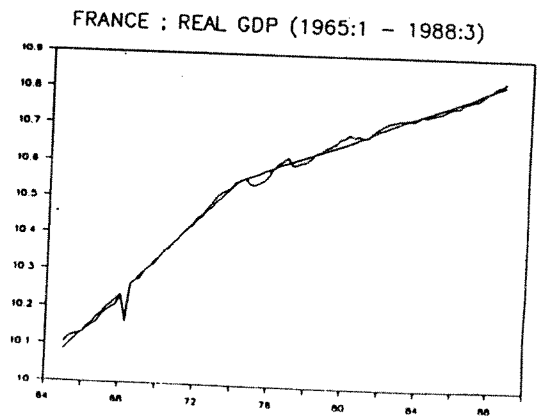

Figure 4

GERMANY ; REAL GNP (1960:1 – 1986:2)

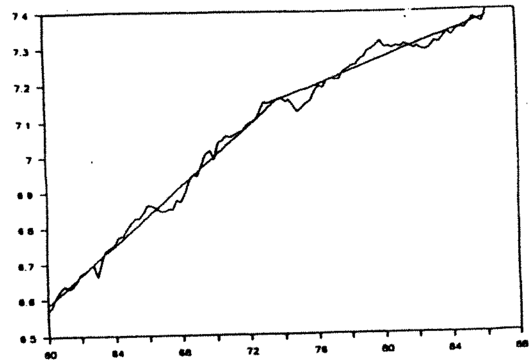

Figure 5

ITALY ; REAL GDP (1960:1 – 1985:1)

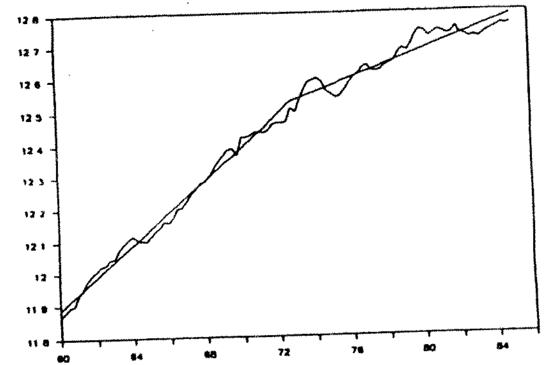

Figure 6

UK ; REAL GDP (1957:1 – 1986:2)

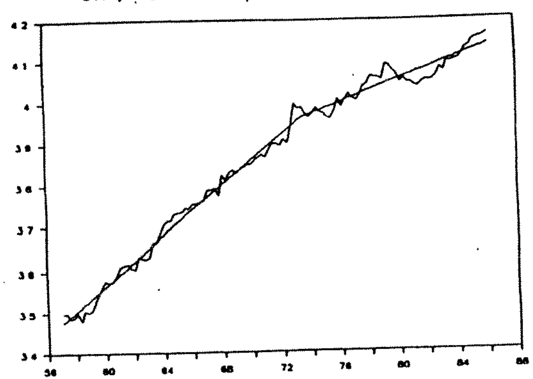

Figure 7

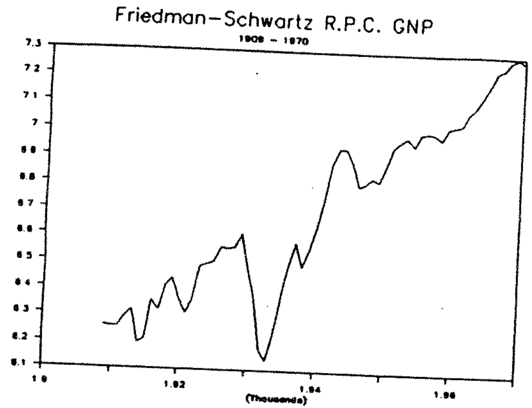

Figure 8

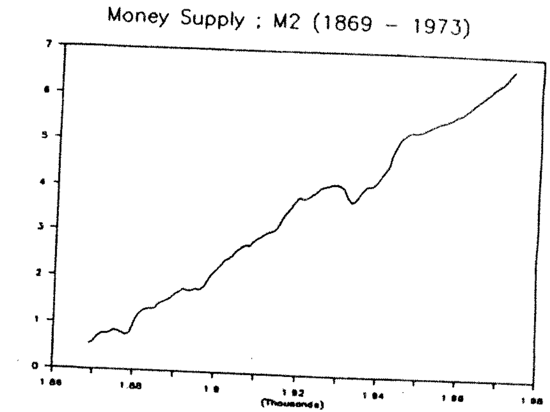

Figure 9

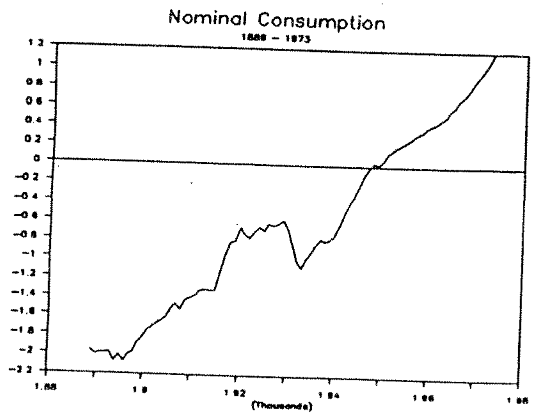

Figure 10

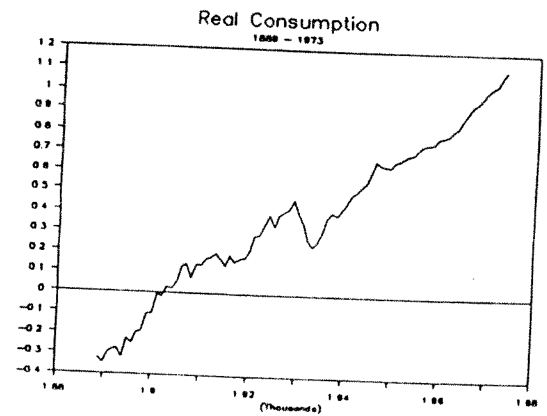

Figure 11

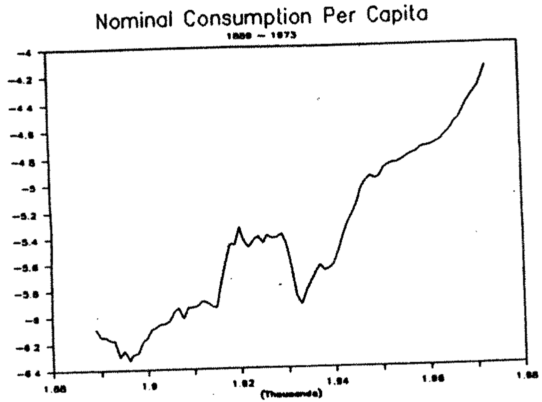

Figure 12

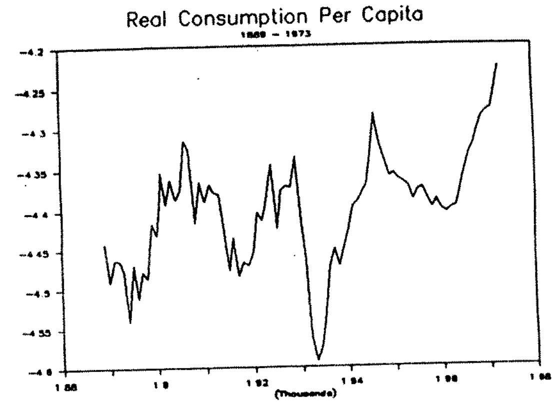

Figure 13

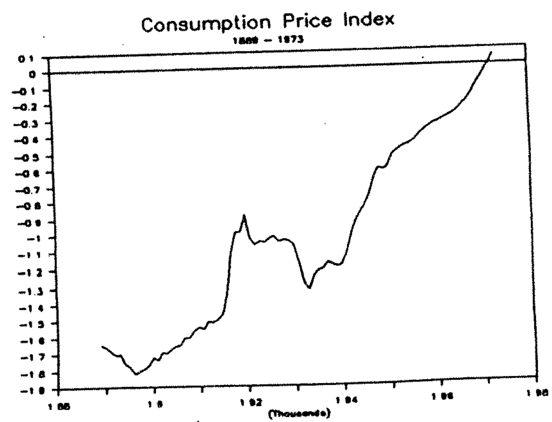

Figure 14

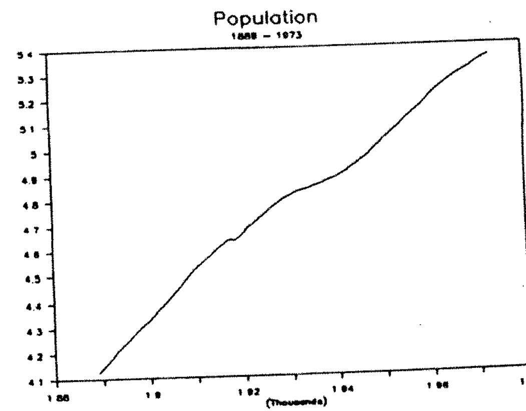

Figure 15

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