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No.423

WARWICK ECONOMIC RESEARCH PAPERS

DEPARTMENT OF ECONOMICS

UNIVERSITY OF WARWICK
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1. Introduction

In a series of papers (Crafts, Leybourne & Mills, 1989a, 1990, 1991), we have analysed British industrial production over the period 1700 to 1913 using various time series approaches to decompose an index of this series into its trend and cyclical components.

The publication of these papers, along with related work by Crafts (1983, 1987, 1989) and Harley (1982), has prompted numerous responses, in which our work is criticised in three basic areas. The first two criticisms, that our estimates of industrial production are incorrect and that our view of the industrial revolution is misconceived, have been the subject of detailed responses and rebuttals by Crafts & Harley (1992) and Harley & Crafts (1994). It is the third area of criticism, the appropriate way of modelling the process generating industrial production, that forms the basis for the present paper. The two major critiques of our econometric work are Newbold & Agiakloglou (1991) and Greasley & Oxley (1994), to which we have provided detailed responses in Crafts & Mills (1992, 1994a). We do not repeat these debates in detail here: rather we attempt, very much in the spirit of a progressive research strategy, to extend and develop the existing models of industrial production. This is in response both to our critics and to recent developments in time series econometrics, and enables us to provide a framework that encompasses previous models and which allows us to present further evidence in support of our views on the timing and extent of the industrial revolution and on the presence or otherwise of the climacteric of the late nineteenth century (see Crafts, Leybourne & Mills, 1989b).

Section 2 thus outlines the approach that we have taken in our initial work and briefly discusses the recent critiques on this that have been made by Newbold & Agiakloglou (1991) and Greasley & Oxley (1994). The general theme of these critiques is that the process generating industrial production has undergone various shifts during the period and it is this insight that acts as the point of departure for our present analysis. Employing a slightly revised index of industrial production, Section 3 investigates the question of structural breaks using, *inter alia*, some recently developed unit root tests. These tests strongly suggest that both difference stationary and segmented linear trend models are inadequate processes for explaining the behaviour of industrial production and this leads us in Section 4 to investigate the usefulness of an alternative model, that of a segmented quadratic trend plus a shifting cyclical component. This does indeed provide a satisfactory representation of the data and explains the models of our critics. An interpretation of the results in terms of the historiography of British economic growth is provided in the concluding Section 5.

2. Trends and Cycles in Industrial Production Revisited

The approach that we have favoured in our previous work is that of the *basic structural model* (BSM: see Harvey, 1985, 1989), in which the logarithm of industrial production, y_t say, is additively decomposed into a trend, μ_t, and a cycle, ψ_t,

$$y_t = \mu_t + \psi_t \tag{1}$$

such that the trend is modelled as

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t \tag{2a}$$

$$\beta_t = \beta_{t-1} + \xi_t \tag{2b}$$

The cycle may be modelled as an AR(2) process,

$$\psi_t = \rho_1\psi_{t-1} + \rho_2\psi_{t-2} + \omega_t \tag{2c}$$

This is, of course, stationary if $\rho_1+\rho_2<1$, and displays pseudo-cyclical behaviour if $\rho_1^2+4\rho_2<0$. The period of the cycle may be calculated as

$$p = 2\pi/\cos^{-1}(|\rho_1|/2|\rho_2|^{1/2})$$

Alternatively the cyclical component may be modelled as an explicit sinusoidal process:

$$\begin{pmatrix}\psi_t\\ \psi_t^*\end{pmatrix} = \rho\begin{pmatrix}\cos\lambda & \sin\lambda\\ -\sin\lambda & \cos\lambda\end{pmatrix}\begin{pmatrix}\psi_{t-1}\\ \psi_{t-1}^*\end{pmatrix} + \begin{pmatrix}\omega_t\\ \omega_t^*\end{pmatrix} \tag{2d}$$

where ψ_t^* appears by construction and where the parameters $0\le\lambda\le\pi$ and $0\le\rho\le1$ have direct interpretations as the frequency of the cycle ($\lambda=2\pi/p$) and the damping factor on the amplitude, respectively. This process is stationary if $\rho<1$.

The errors η_t, ξ_t, ω_t and ω_t^* are independent white noise disturbances with variances σ_η^2, σ_ξ^2, σ_ω^2 and σ_ω^{*2} (usually assumed to equal σ_ω^2 in estimation). The trend component is thus modelled as a *stochastic* linear trend, which would collapse to a *deterministic* linear trend if the variances σ_η^2 and σ_ξ^2 were both zero, and to a random walk with drift if only $\sigma_\xi^2=0$. If $\sigma_\omega^2=0$ then the cycle is deterministic.

In Crafts, Leybourne & Mills (1989a, henceforth CLM), we fitted this model (using the autoregressive cyclical process (2c)) to the full sample of industrial production data and to two subsamples, 1700-1783 and 1815-1913, finding that the parameter estimates altered substantially across samples. For the full sample all variances were significantly different from zero, whereas for the earlier 'pre-take-off' period σ_ξ^2 was zero, implying that trend growth was constant, and ρ_1 and ρ_2 were insignificantly different from zero, so that there was no cycle in the data. For the post-Napoleonic period, however, $\sigma_\eta^2=0$, so that permanent changes in the level of the series were brought about entirely by changes in trend, and the positive estimate for σ_ω^2 combined with significant estimates of the ρ's implied a stochastic cycle with a period of approximately 7.5 years.

Since it is well known that the BSM implies that y_t has a (restricted) ARIMA(2,2,4) representation and, if $\sigma_\xi^2=0$, an ARIMA(2,1,3) representation, we also undertook a Box-Jenkins analysis, finding that for the full sample conventional identification led to an ARIMA(0,1,1) model. Since the first-order autocorrelation of ∇y_t is negative, this is consistent with the structural model (1) with

$$\mu_t = \mu_{t-1} + \beta + \eta_t$$

and $\psi_t = \omega_t$: i.e. constant trend growth and no cyclical component. This was also consistent with standard unit root tests, which for all samples could not reject the difference stationary null in favour of a (linear) trend stationary alternative.

However, the ARIMA(2,2,4) model, although fitting substantially worse than the ARIMA(0,1,1), was able to model the overall autocorrelation structure of the series much more satisfactorily, for Ljung-Box portmanteau statistics calculated over 36 lags were 51.7 for the latter (significant at a marginal significance level of less than 3%) but only 27.3 for the former. This, we argued, was consistent with Harvey's (1985) argument that structural time series models are not intended as parsimonious representations of the underlying data generation process but, in the present context, aim to present the historiography of the series in terms of a decomposition that is of interest to economists and economic historians. We might also profitably recall Cochrane's (1988) argument that ARIMA models are designed to fit the low order autocorrelations of the data as closely as possible at the expense of accurately fitting the higher orders, an important consideration for short-term forecasting but not necessarily so in the present context.

Notwithstanding these arguments, Newbold & Agiakloglou (1991) took issue with our use of BSMs, preferring themselves to remain within the ARIMA framework. They demonstrated that unit root tests on ∇y_t conclusively reject the presence of the two unit roots implied by the BSM for the different subsamples and that, within this framework, the case for evolving (non-constant) growth rates and a cyclical component in the post-1815 period was rather weak. Our defence of BSMs within the present context is already contained in Crafts & Mills (1992), but what is clear from both CLM and Newbold & Agiakloglou is that the assumption of a constant model structure throughout the entire sample is very difficult to maintain.

Greasley & Oxley (1994) also focus on the issue of structural stability: on considering three subsamples, 1700-1780, 1781-1851 and 1852-1913, they find that unit root tests reject the null of difference stationarity in favour of (linear) trend stationarity for the first and third subsamples, but cannot for the second. Their interpretation of these findings, that the implied persistence of innovations to industrial production during the period from 1781 to 1851 marks this out as a 'distinctive macroeconomic epoch', has been criticised in Crafts & Mills (1994a), but what is important here is Greasley & Oxley's additional evidence of instability in the process generating industrial production, an instability that obviously requires further investigation and which acts as a point of departure for our own analysis.

3. Structural Breaks in Industrial Production

Crafts & Harley (1992) present some revisions to what has become known as the CLM index. In what follows we thus use their Revised Buest Guess index of industrial production, described in Appendix 3 and reported in Table A3.1 of Crafts & Harley (1992). The logarithms of the index are plotted in Figure 1.

Given the evidence presented above that the index may contain structural breaks, the BSM was estimated over both the full sample 1700-1913, and various subsamples suggested by the literature. These estimates are shown in Table 1, from which it would appear that σ_{ξ}^{2} was close to zero for subsamples containing the middle part of the nineteenth century, while σ_{η}^{2} was close to zero for the latter part of the sample, this being the only time when σ_{ω}^{2} was positive and there was a significant cycle, its estimated period being approximately 8 years. Although for all other subsamples $\sigma_{\omega}^{2}=0$, implying a deterministic cycle, the estimated periods were all in excess of 20 years and examination of the sample autocorrelations for ∇y_t showed cut-offs at lag 2 rather than cyclical behaviour (cf. Harvey, 1985: all estimations were performed using the sinusoidal process (2d) for the cycle ψ_t).

It is clear from these estimates that a single BSM fitted over the full sample is unlikely to provide a completely satisfactory fit and further evidence that this is the case can be provided from another source. The full sample estimates show that although σ_{η}^{2} is reliably positive, σ_{ξ}^{2} is very small and could be taken as being essentially zero. Although $\sigma_{\omega}^{2}=0$, for the reasons given above, there is no cycle in the error component.

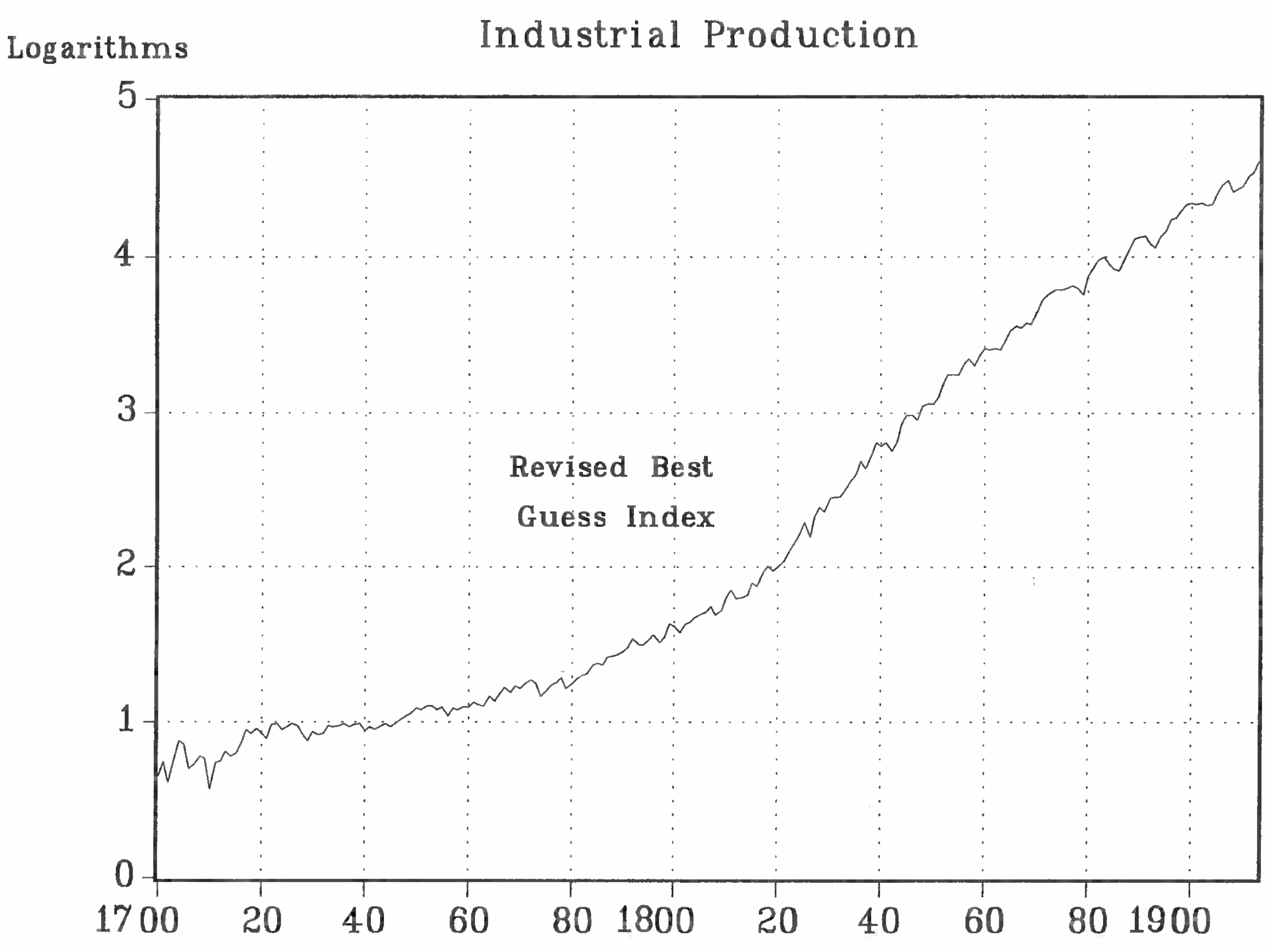

Industrial Production
Logarithms
5
4
3
2
1
0
Revised Best
Guess Index
1700
20
40
60
80
1800
20
40
60
80
1900

Figure 1

Table 1

BSM estimates for Subperiods

	σ^2_{η} ($\times 10^{-3}$)	σ^2_{ξ} ($\times 10^{-5}$)	σ^2_{ω} ($\times 10^{-4}$)
1700-1913	2.26 (10.03)	0.55 (0)	0
1700-1780	3.04 (5.98)	2.70 (0.96)	0
1781-1829	1.90 (4.47)	4.50 (0.94)	0
1781-1850	2.11 (5.76)	0.18 (0)	0
1813-1913	1.91 (6.61)	2.66 (1.19)	0
1830-1913	1.66 (5.87)	4.31 (1.26)	0
1830-1872	1.83 (4.40)	0.77 (0.72)	0
1851-1913	0 (0)	0.10 (0)	6.67 (4.85)
1873-1913	0 (0)	8.03 (1.52)	4.51 (3.37)

t-ratios in parentheses.

Thus, with $\sigma^2_{\xi}=0$ and ψ_t modelled as a noncyclical AR(1) process, it is easy to show that ∇y_t must follow an ARMA(1,2) process with drift. Fitting such a model obtains

$$\nabla y_t = \underset{(.001)}{.018} - \underset{(.125)}{.691}\nabla y_{t-1} + a_t + \underset{(.129)}{.570}a_{t-1} - \underset{(.115)}{.214}a_{t-2}, \quad \hat{\sigma}_a = 4.64\% \qquad (3)$$

Although this model appears to be quite satisfactory on the basis of standard diagnostic checks, on further analysis we find that ∇y_t contains a significant time trend, for including t as an additional regressor in (3) yields a t-statistic of 3.23. This implies that ∇y_t is nonstationary; $\nabla^2 y_t$ is stationary though, albeit around a nonzero mean and with a unit root in its moving average part. The presence of such a nonzero mean has the important implication that y_t cannot be regarded as being generated from a BSM. Moreover, the presence of a linear trend in ∇y_t implies a quadratic trend in y_t. Newbold & Agiakloglou (1991) are thus incorrect in claiming that the rejection of two unit roots in favour of there being only one precludes there being evolving trend growth rates: their unit root tests excluded a time trend and led them to conclude that ∇y_t was stationary around a constant mean rather than, as we have found, around a linear trend.

Given these twin findings of structural instability in the BSM and of a full sample ARMA process for ∇y_t that does not admit a BSM representation but contains a linear trend, it would seem fruitful to consider an alternative approach to modelling trends and cycles in the Revised Best Guess index.

An obvious possibility is that the presence of a linear trend in the full sample model for ∇y_t is compensation for the trend component μ_t being a segmented linear function: a 'changing growth' model in the terminology of Perron (1989). The timing of such a shift in trend, which is here being 'approximated' by a quadratic in t buried in nonstationary noise, will not be known precisely, so that Perron's analysis, which assumes a single, known, break point, will not be appropriate. Banerjee, Lumsdaine & Stock (1992), however, have developed a set of recursive and sequential tests for detecting such a break when its timing is unknown.

Thus we may consider the model

$$y_t = \mu_0 + \mu_1 t + \mu_2 \tau_t(k) + \alpha y_{t-1} + \sum_{i=1}^{p} \beta_i \nabla y_{t-i} + a_t \qquad (4)$$

for $t=1,\ldots,T$, where p is known, a_t is a martingale difference sequence, and $\tau_t(k)=(t-k)\cdot 1(t>k)$, where $1(\cdot)$ is the indicator function. $\tau_t(k)$ is thus a deterministic regressor capturing the possibility of a shift in the trend at time k.

Banerjee, Lumsdaine & Stock (1992), henceforth BLS, consider various tests for structural breaks in this model. With μ_2 assumed to be zero, (4) becomes a standard Dickey-Fuller (1979) regression, so that when it is estimated by OLS using the full sample $t=1,\ldots,T$ the t-statistic testing $\alpha=1$, $\hat{t}_{DF}$, is their $\hat{\tau}_\tau$ test for a unit root against a linear trend stationary alternative. Such a test is well known to be incapable of rejecting the unit root null when the alternative is that of a shifting trend. The test can, however, be computed recursively, i.e. using subsamples $t=1,\ldots,k$, for $k=k_0,\ldots,T$, where $k_0=[\delta_0 T]$ is a start up value, thus leading to the set of statistics $\hat{t}_{DF}(k/T)$ (noting that $\hat{t}_{DF}(1)=\hat{t}_{DF}$, the full sample statistic). BLS show that functions of this set of statistics can have power against trend-shift alternatives of the form (4).

The statistics can also be computed from rolling regressions, i.e. using subsamples that are a constant fraction δ_0 of the full sample, rolling through the sample. Thus the regression (4) is estimated over the subsamples $k-k_0+1,\ldots,k$, for $k=k_0,\ldots,T$, yielding the set of statistics $\tilde{t}_{DF}(k/T)$.

BLS provide asymptotic critical values and examine the size and power of various functions of these sets of unit root statistics. Their simulations suggest that the most useful statistics for examining the possibility of shifting roots and shifting trends are the maximal and minimal Dickey-Fuller statistics $\hat{t}_{DF}^{max}=\max_k \hat{t}_{DF}(k/T)$ and $\hat{t}_{DF}^{min}=\min_k \hat{t}_{DF}(k/T)$, with similar definitions for $\tilde{t}_{DF}^{max}$ and $\tilde{t}_{DF}^{min}$.

With μ_2 allowed to be nonzero, equation (4) corresponds to Perron's (1989) 'changing growth' model in which there is a segmented linear trend with the break at time k. We may now consider sequential Dickey-Fuller statistics calculated over the full sample but which allow k to move through the range $k_0 \leq k \leq T-k_0$, denoted $t_{DF}^*(k/T)$. BLS focus attention on t_{DF}^{min*} and also on two further statistics: the maximum of the sequential F statistics, F^{max}, testing the hypothesis $\mu_2=0$, and the associated Dickey-Fuller statistic evaluated at the value of k, $\hat{k}$, that maximises F, $t_{DF}^*(\hat{k})$.

A related sequential statistic considered by BLS is the Quandt (1960) likelihood-ratio (LR) statistic, which tests for a break in any or all of the coefficients. This entails estimating (4), with $\mu_2=0$, over pairs of subsamples $1,\ldots,k$ and $k+1,\ldots,T$, for $k_0 \leq k \leq T-k_0$, computing Quandt's LR statistic for each each break point and considering the maximum of these, Q_{LR}.

These statistics were computed for the Revised Best Guess Index after setting the number of lags of ∇y_t entering the regression (4) at $p=3$, the number that were found to be significant in the full sample regression. A variety of settings of the 'trimming' parameter δ_0 were investigated, and the set of statistics are reported in Table 2. Precise inferences are difficult because BLS only report critical values for a limited range of T and δ_0 values. However, the recursive and sequential Dickey-Fuller statistics are constant across δ_0 and are clearly insignificant. The rolling statistics do vary and, although $\tilde{t}^{max}_{DF}$ is certainly insignificant, $\tilde{t}^{min}_{DF}$ appears to be significant for low values of δ_0, as does Q_{LR}. It is of interest to note that F^{max}, $\tilde{t}^{min}_{DF}$ and Q_{LR} occur at very different years: 1775, 1827 and 1870 respectively.

These various pieces of evidence suggest that, although there is certainly evidence of structural instability, it is unlikely to be characterised by a single shift in a linear trend function.

Table 2

Recursive, Rolling and Sequential Test Statistics for Shifting Roots and Trends

	Recursive		Rolling		Sequential			
δ_0	$\hat{t}^{max}_{DF}$	$\hat{t}^{min}_{DF}$	$\tilde{t}^{max}_{DF}$	$\tilde{t}^{min}_{DF}$	t^{min*}_{DF}	F^{max}	$t^*_{DF}(\hat{k})$	Q_{LR}
$\frac{1}{6}$	2.12	-2.67	1.23	-5.21	-3.45	9.07	-3.45	41.32
$\frac{1}{5}$	2.12	-2.67	1.43	-5.22	-3.45	9.07	-3.45	41.32
$\frac{1}{4}$	2.12	-2.67	1.01	-4.10	-3.45	9.07	-3.45	32.33
$\frac{1}{3}$	2.12	-2.67	1.14	-3.74	-3.45	9.07	-3.45	25.63

4. A Segmented Quadratic Trend Model

The likely presence of more than one structural break, plus the visual evidence from Figure 1 of nonlinearity in the evolution of the Revised Best Guess index, suggests that a segmented polynomial trend might profitably be considered. Segmented (or piecewise or grafted) polynomials have a long history in fitting smooth trend functions to time series and have a firm mathematical foundation (see Fuller, 1969, 1976). They have been used in a similar context by Hausman & Watts (1980) and also by Crafts & Mills (1994b) in fitting trends to real wages. Although cubics are popular in curve fitting, we felt that, given the evidence of a time trend in ∇y_t, a segmented quadratic was the appropriate function to consider.

The received historiography plus the evidence from Section 3 suggests that major changes in the growth rate of the index took place during the years 1765-1785, 1815-1835 and 1855-1875. The first of these intervals is quite generally thought to mark the start of the 'industrial revolution' and an associated upturn in the rate of growth; authors who write in these terms include Fisher (1982), Hoffman (1955), Hudson (1992), McCloskey (1981), Mokyr (1993) and Rostow (1960). The second interval takes in suggested terminal dates for the industrial revolution from several writers, including Hudson and Mokyr, when growth might be held to have entered a mature phase and acceleration of the growth rate would be expected to have ceased. Others, including Fisher, Hoffman, McCloskey and Rostow, have placed this point later, argiung that it was located in the third quarter of the nineteenth century, while a more pessimistic literature deals with a climacteric in growth arriving in the 1870s (Saul, 1985). Our third proposed interval captures these alternative views.

Given this argument, we thus chose to model the index as having three breaks, denoted T_1, T_2 and T_3, respectively, so that the following trend component was considered:

$$\mu_t = \mu_0 + \mu_1 t + \mu_2 t^2 + \sum_{i=1}^{3} \vartheta_i \tau_{it}^2 \tag{5}$$

where $\tau_{it}^2=(t-T_i)^2 \cdot I(t>T_i)$ and $1764<T_1<1786$, $1814<T_2<1836$ and $1854<T_3<1876$. The times of the breaks were determined by minimising the residual sum of squares from the regression $y_t=\mu_t+\psi_t$, with ψ_t assumed to be white noise (i.e. $\psi_t=\omega_t$), for all possible combinations of T_1, T_2 and T_3 (for a justification of this approach, see Gallant & Fuller, 1973). The minimising combination was found to be T_1=1776, T_2=1834 and T_3=1874, which are relatively close to the breaks points informally suggested by the analysis of Section 3. Indeed, the fit of the model with break points at 1775, 1827 and 1870 was not drastically worse, the residual standard error being 5.10% as opposed to 4.90%.

Table 3 presents the OLS estimates from fitting this model. In line with the approach of Campbell & Perron (1991), a unit root test on the fitted residuals $\hat{\psi}_t$ produces a test statistic of -8.10. Approximate critical values, conditional on the estimated break points, were obtained via Monte Carlo simulation, from which it was found that the 1% critical value was -5.23. Even allowing for pre-testing bias, such a large value of the test statistic

Table 3

Segmented Quadratic Model

	OLS		GLS	
	Est	t-rat	Est	t-rat
μ_0	0.730	50.20	0.725	41.70
μ_1	0.0064	9.47	0.0070	25.23
$\mu_2 \times 10^{-4}$	0.006	0.99	-	-
$\vartheta_1 \times 10^{-4}$	2.499	20.58	2.628	35.74
$\vartheta_2 \times 10^{-4}$	-4.395	-23.95	-4.550	-16.85
$\vartheta_3 \times 10^{-4}$	1.422	3.66	1.675	2.62
$\hat{\rho}$	-		0.520	8.76
$\hat{\sigma}_\omega$	4.90%		4.18%	
dw	0.96		1.92	
t_{DF}	-8.10		-	

points strongly to rejecting the hypothesis of a unit root, and this is confirmed by restimating the equation by GLS assuming an AR(1) process for ψ_t, the estimates of which are also shown in Table 3. Extending the order of the autoregressive error process did not produce significant coefficients, thus confirming that, for the full sample, there is no cyclical behaviour about trend. In this GLS estimation the insignificant variable t^2 was omitted: this has the effect of forcing the first segment to be linear, but the significance of the τ_i^2 terms confirms that the remaining two segments are both quadratic in t.

However, as we have seen from the subsample BSM estimates, there is strong evidence of a cycle in the later years of the sample. We thus investigated the possibility of there being a shifting autoregressive error process, with the same break points being used as above. No breaks in the process were found at 1776 or 1834, but an important break was isolated at 1874, leading to the model

$$\tilde{\psi}_t = \begin{cases} 0.533\tilde{\psi}_{t-1} + a_t & t \le 1874 \\ 0.533\tilde{\psi}_{t-1} - 0.486\tilde{\psi}_{t-3} + a_t & t > 1874 \end{cases}, \quad \hat{\sigma}_a = 3.93\% \qquad (6)$$

where $\tilde{\psi}_t$ are the residuals obtained as $y_t - \tilde{\mu}_t$, $\tilde{\mu}_t$ being the trend fitted from the GLS estimates. The complex roots in the post-1874 AR(3) process yield a period of 9 years, which accords well with the estimated cycle from the BSM model fitted over a similar subsample.

The segmented trend $\tilde{\mu}_t$ is shown superimposed on y_t in Figure 2, while the cyclical component is shown in Figure 3. Trend growth, $\nabla\tilde{\mu}_t$, is shown in Figure 4 and is a very smooth curve, with the major changes in curvature occuring in the early part of the nineteenth century. Trend growth presents a pattern very similar to those reported in Crafts, Leybourne & Mills (1989a, 1990, 1991) and Crafts & Harley (1992), there being modest and constant growth of 0.65% per annum for most of the eighteenth century, an acceleration between 1776 and 1834, where a maximum of approximately 3.7% per annum is reached, and a deceleration until 1874, from whence growth gently declines to 2% per annum by 1913.

It should be recalled that Feinstein, Matthews & Odling-Smee (1982) have suggested that there was an Edwardian climacteric in

Industrial Production

Actual and Segmented Trend

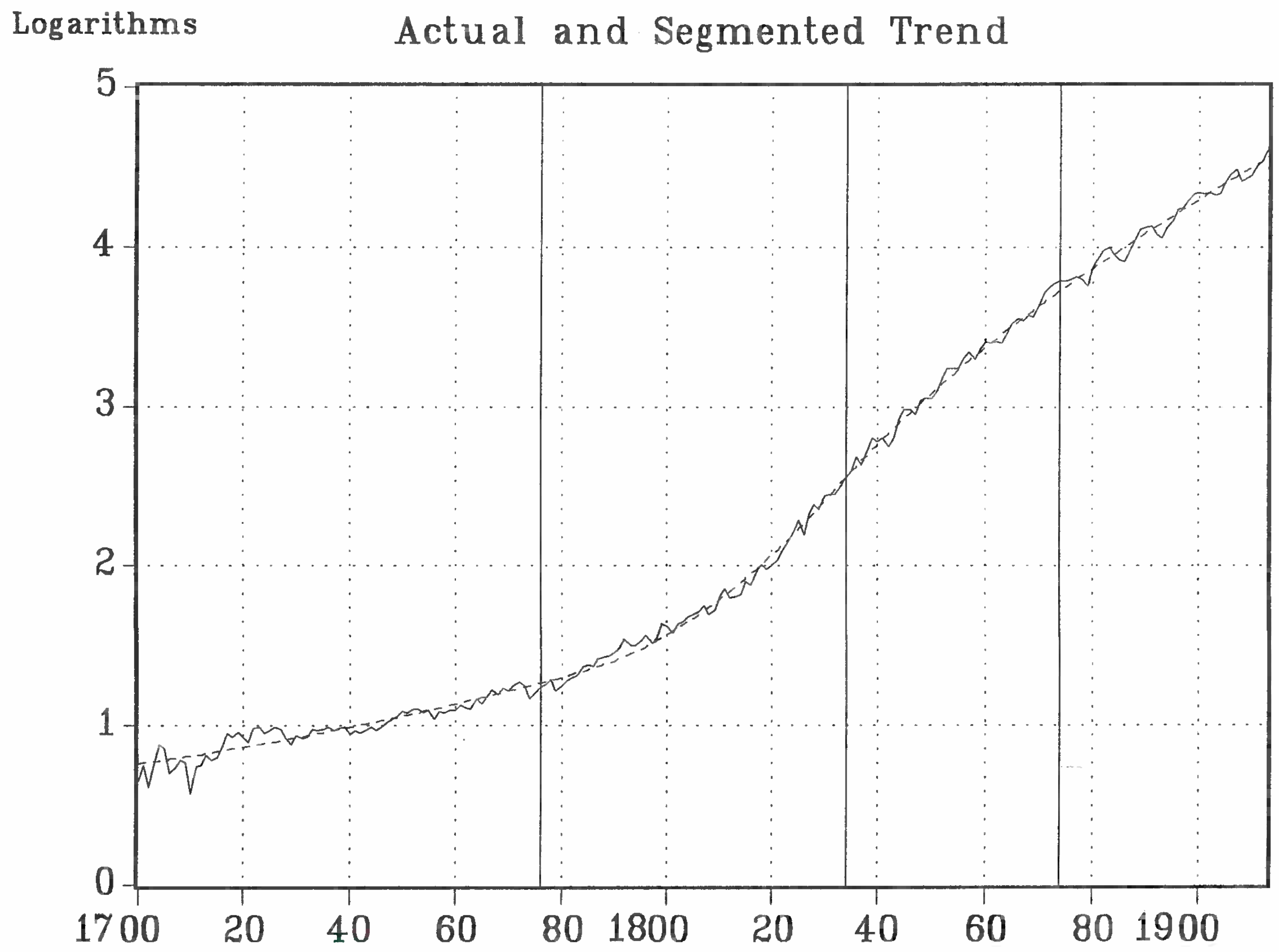


Figure 2

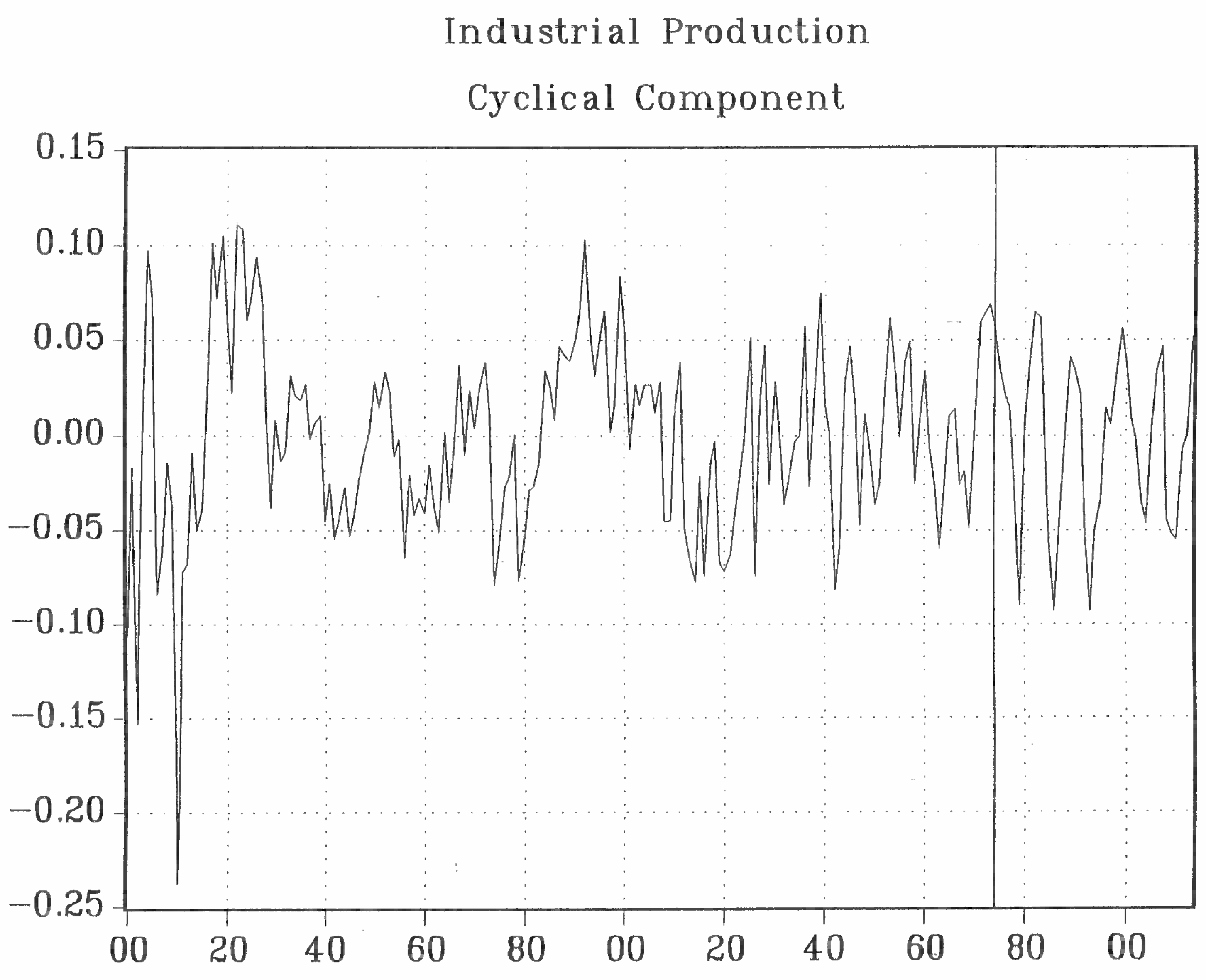
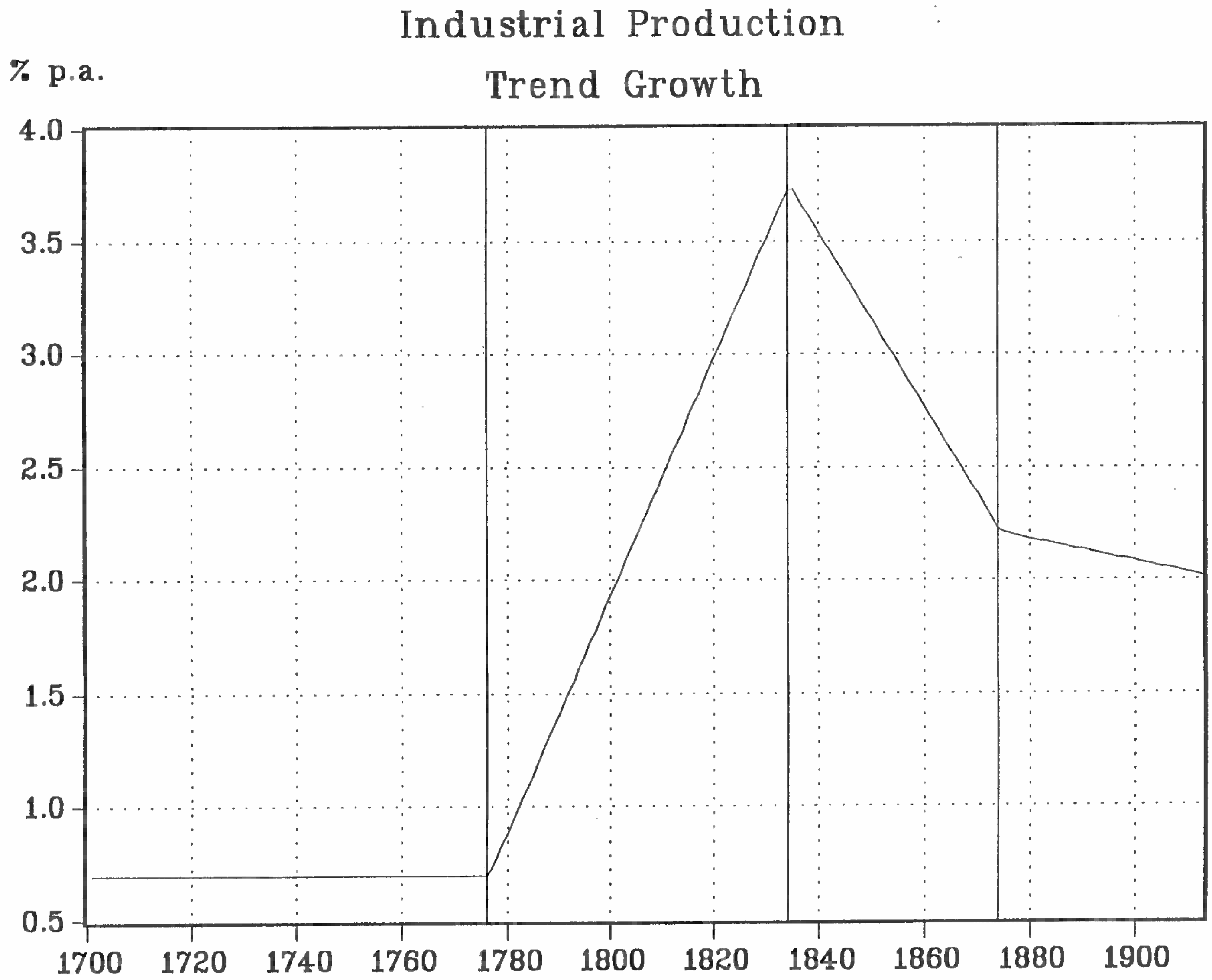

Industrial Production
Cyclical Component
0.15
0.10
0.05
0.00
−0.05
−0.10
−0.15
−0.20
−0.25
00
20
40
60
80
00
20
40
60
80
00

FIGURE 3



Industrial Production
Trend Growth
% p.a.
4.0
3.5
3.0
2.5
2.0
1.5
1.0
0.5
1700
1720
1740
1760
1780
1800
1820
1840
1860
1880
1900

Figure 4

output growth. Although we have previously rejected the hypothesis of a climacteric after 1899 (Crafts, Leybourne & Mills, 1989b), we nevertheless estimated the fitted model (5) with an additional break in the interval from 1895 to 1905 and failed to find any significant break point, the t-statistics on the additional variable always being greater than -0.5.

The above pattern of trend growth may be compared with that implied by the methodology of Greasley & Oxley (1994). Greasley & Oxley report only the Dickey-Fuller test statistics for the subsamples 1700-1780, 1781-1850 and 1851-1913, but it is of some interest to report the models underlying their calculation. Using our framework, their models can be written as

$$y_t = \mu_0 + \mu_1 t + \psi_t$$

and

$$\psi_t = \sum_{i=1}^{3} \pi_i \psi_{t-i} + \omega_t$$

Table 4 presents estimates of this model for the three subsamples, reporting Dickey-Fuller statistics for the unit root hypothesis $\pi_1+\pi_2+\pi_3=1$. The linear trend stationary models for the first and third subsamples yield constant trend growth of 0.065% per annum from 1700 to 1780 and 2.24% per annum from 1851 onwards. Imposing a unit root on the model for the 1781-1850 period reveals that an MA(1) model provides an adequate fit to ∇y_t, i.e. $\nabla y_t=\theta_0+a_t-\theta_1 a_{t-1}$, where $\hat{\theta}_0=0.0260$ and $\hat{\theta}_1=0.444$.

A trend component can be computed from such a model using the technique of signal extraction on the assumption that μ_t follows a random walk with drift, i.e. $\mu_t=\theta_0+\mu_{t-1}+\eta_t$ (the special case of the BSM with $\sigma_\xi^2=0$), and that the cycle ψ_t is white noise. In this

Table 4

Estimates of Model (1)

	1700-1780	1781-1850	1851-1913
μ_0	0.738 (0.022)	-2.645 (2.127)	-0.190 (0.092)
μ_1	0.0065 (0.0005)	0.0373 (0.0119)	0.0224 (0.0005)
π_1	0.493 (0.098)	0.573 (0.092)	0.923 (0.118)
π_2	-	-	-0.382 (0.116)
π_3	-	0.375 (0.092)	-
$\hat{\sigma}_\omega$	4.80%	4.03%	3.28%
$\sum\pi_i$	0.493	0.948	0.541
t_π	-5.20	-1.07	-5.09

$\sum\pi_i=\pi_1+\pi_2+\pi_3$; t_π tests the hypothesis $\sum\pi_i=1$.
Standard errors are shown in parentheses.

case the trend is estimated as

$$\hat{\mu}_t = \frac{(1-\theta_1)^2}{1-\theta_1^2} \sum_{i=-\infty}^{\infty} \theta_1^{|i|} y_{t-i}$$

Truncating this computation at $i=8$ yields a trend component that, as implied above, has a mean of 2.60% per annum, but there is considerable variability about this mean value, the standard deviation being 1.47% with a maximum value of 6.20% and a minimum of -0.24%. Modelling industrial production as an $I(1)$ process over the period 1781-1850, so that innovations have a permanent effect on the series, does not lead to a trend component having the degree of smoothness that we would commonly wish for in such a situation.

5. Implications for the Historiography of British Economic Growth

Obviously the details of our modelling of industrial output growth has been determined by econometric criteria. At the same time, our strategy has been informed by the historiography of British industrialisation and we believe that our results reinforce some important aspects of economic historians' accounts of eighteenth and nineteenth century growth. This section sets out these features of our research.

The most important insights into growth which provide the background to our approach come from the literature on the economic history of technological change, particularly as organised and interpreted by Mokyr (1990, 1993). Mokyr suggests that "a technological definition of the industrial revolution is a clustering of macroinventions leading to an acceleration in microinventions" (1993, p. 22).

'Macroinventions' are defined as "those inventions in which a radical new idea, without clear precedent, emerges more or less ab nihilo" (Mokyr, 1990, p. 13). 'Microinventions', which account for a very high fraction of all productivity increases, are influenced by economic factors and come about notably through learning by doing and learning by using. The process of microinvention within any particular technology is subject to diminishing returns. Without macroinventions, productivity growth would eventually dry up, yet macroinventions cannot be predicted and often result from strokes of luck or genius (Mokyr, 1990, p. 13). Macroinventions of the eighteenth century were crucial to the growth of the Victorian staples, including cotton, coal and iron & steel.

Several aspects of this view are reflected both in our modelling strategy and in our results. These include the following:

(i) The growth process was disturbed by technological surprises which could change the trend rate of growth.

(ii) We would expect that productivity and output growth would not be fully endogenous in the sense of Rebelo (1991). Instead, major technological shocks would be followed by rising, but then falling, rates of output growth.

(iii) In general, there will be no strong priors about the precise path of the growth process. Even dating of famous macroinventions does not help since their impact on growth will not be immediate and might be quite long delayed.

We therefore wanted to model changing trend growth allowing break points not to be precisely specified and to subject any apparent failure to reject the unit root hypothesis to close scrutiny.

In the present case, it is fairly straightforward to sketch out an account which puts historical flesh on the econometric bones. First, there is no doubt that contemporaries did not foresee the industrial revolution nor did they immediately appreciate the significance of what, with hindsight, were macroinventions with major consequences. For example, Adam Smith did not understand the importance of the inventions by James Watt, James Hargreaves and Richard Arkwright and "did not suppose that England was about to embark on a period of unprecedented gain in output per head" (Wrigley, 1987, p. 21).

Second, the full impact of technological developments built up over time. For example, the steam engine (invented in 1765) supplied less than 35,000 HP in 1800 but 2 million in 1870 (Kanefsky, 1979, pp. 373-5). The importance of learning by doing and the relatively lengthy diffusion of the improved versions of famous inventions have been clearly laid out for both the cotton (Chapman, 1972, David, 1975, and von Tunzleman, 1978) and iron industries (Allen, 1983, and Hyde, 1977).

Third, the experience of slowdown following the erosion of growth possibilities of tecnological breakthroughs is a familiar one in nineteenth century British economic history. It is, in fact, precisely the vision embodied in the seminal paper of Phelps-Brown & Handfield-Jones (1952), which first put forward the idea of a late nineteenth century climacteric. The slowing down of total factor productivity growth based on the classic industrial revolution technologies in both cotton and iron is well known (Lazonick & Mass, 1984, and McCloskey, 1973). We have shown in an earlier paper that slowdown in industrial growth was relatively pronounced in the staple industries (Crafts, Leybourne & Mills, 1991, p. 140). For the whole economy, a common feature of growth accounting exercises for the nineteenth century is a pattern of steadily rising and then declining total factor productivity growth (Crafts, 1993).

In the end, the estimates graphed in Figure 4 are consistent both with this account of technological change and also with the description of trend growth in our earlier work, as we noted in Section 4. We still find a long period of rising trend growth to the mid-1830s and we still see the beginnings of declining trend growth starting relatively early. We also continue to find much stronger trend growth in late Victorian Britain than in the pre-industrial revolution phase of Smithian growth.

Finally, as our discussion of technological change anticipated, we reject the hypothesis of a unit root in industrial output which would be predicted by endogenous growth models of the Rebelo (1991) type. We prefer to retain the notion of exogenous technological shocks rather than to work with models which seek completely to endogenise productivity change.

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