Structural Breaks and Convergence in Output Growth in the EU

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Abstract. Convergence is defined for a multivariate time-series model of output with breaks in intercepts and in time trends. Using OECD quarterly data on output from 1980, the convergence hypothesis is tested across seven European economies, Belgium, Finland, France, Italy, the Netherlands, Spain and the UK. On the strictest definition, the hypothesis of convergence of output can be rejected but, with a weaker definition, there is some evidence of convergence for the five countries Belgium, Finland, France, the Netherlands and the UK. The data is consistent with a model in which each country's trend output is related to a common European stochastic trend. This trend output is estimated and graphed for each country.

Key words: Convergence, Output Growth, Permanent Stochastic Trends.

JEL Classification: C32; C51.

1. Introduction

This paper is concerned with the convergence of long run trends in output in European economies. We consider a number of time-series definitions of convergence and demonstrate how they can be tested as restrictions on the parameters which generate the stochastic trend in GDP, when that trend is taken as the permanent component in a multivariate Beveridge-Nelson [5, (1981)] decomposition. As recent research (e.g., Strazicich, Lee and Day [18, (2001)] and Li and Papell [14, (1999)]) has demonstrated the importance of allowing for structural breaks when testing for convergence, the paper extends the result to the case where there are breaks in mean and breaks in a time trend in the cointegrating equations. The underlying theoretical model with structural breaks is sketched in section 2 and section 3 derives an explicit representation for long run stochastic trends in the presence of such breaks. Section 4 discusses various definitions of convergence and demonstrates the relationship with restrictions on the parameters of the model generating the long run stochastic trends. Section 5 applies the theoretical results to test for convergence in seven European Union countries. Whereas most recent research employs annual data series ending in the mid-nineties (e.g., Bernard and Durlauf [4, (1995)], Strazicich *et al* [18, (2001)] and Li *et al* [14, (1999)]), in this paper we use data for the period 1980Q1 to 2001Q3. Section 6 concludes.

2. The Model

Let $x_t = (x_{1t}, ...x_{pt})'$ be the logarithms of the output series for p European economies. If x_t is $I(1)$ and cointegrated we can write the system in standard VECM form, without structural breaks, as:

$$\triangle x_t = \theta_o + \theta_1 \triangle x_{t-1} + ... + \theta_k \triangle x_{t-k} + \beta\alpha' x_{t-1} + \zeta_t \tag{1}$$

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where $\triangle x_t = x_t - x_{t-1}$, ζ_t is a Gaussian error and α' is the set of cointegrating vectors. If there are structural breaks in either the mean of the VECM and/or the cointegrating relations and shifting time trends in the cointegrating equations [1], then the specification in (1) is inappropriate and may lead to the inference of an incorrect rank for the cointegration space in a multivariate, likelihood ratio test setting. Recent papers by Bai et al [2, 1998], hereafter, BLS, and Johansen et al [13, (2000)] are concerned with identifying structural breaks and testing for cointegration when they are present in a multivariate model.

Suppose there are $q-1$ breaks in the sample with T_1 observations in the first period, T_2 in the second and so on, such that:

$$0 = T_o < T_1 < T_2 \cdots < T_q = T$$

where T is the total number of observations. Johansen et al [13, (2000)] derive a likelihood ratio test for cointegration in the presence of breaks in trend and mean at known points. In general the model can be written:

$$\triangle x_t = \theta_o \Xi_t + \sum_{j=1}^{k} \theta_j \triangle x_{t-j} + \beta(\alpha', \gamma') \begin{pmatrix} x_{t-1} \\ t\Xi_t \end{pmatrix} + \sum_{i=1}^{k+1} \sum_{j=2}^{q} \kappa_{ji} D_{jt-i} + \zeta_t \tag{2}$$

where for $j = 2, ..., q$, $D_{jt} = 1$ for $t = T_{j-1}$and $D_{jt} = 0$ otherwise. The matrix Ξ_t is given by:

$$\Xi_t' = (\Xi_{1t}', ..., \Xi_{qt}')$$

where, for $j = 1, ..., q$, $\Xi_{jt} = 1$ for $T_{j-1} + k + 2 \leqq t \leqq T_j$ and zero otherwise. This specification allows for shifts in the intercepts of both the VECM and cointegrating equations (although they cannot be identified individually), in the term $\theta_o \Xi_t$, and shifts in the trend in the cointegrating equations only, in the term $\gamma' t \Xi_t$. The $\Xi_{jt}s$ are dummies for the effective sample period for each sub-period and the $D_{jt-i}s$ have the effect of eliminating the first $k+1$ residuals of each period from the likelihood, thereby producing the conditional likelihood function given the initial values in each period.

In this paper we restrict the number of breaks to be one, so that $q = 2$, as we only have available twenty-one years of quarterly data, 1980-2001. The procedure of BLS, [2, (1998)] is then available for estimating confidence intervals for break dates in the intercept in a multivariate system.

3. Trend Output

To test for convergence in trend output we use the Beveridge - Nelson [5, (1981)] definition of trend as in Cochrane [7, (1994)], i.e.:

$$x_t^* = \lim_{k \to \infty} E_t(x_{t+k} - k\mu_{\triangle x}) = x_t + \sum_{j=1}^{\infty} E_t(\triangle x_{t+j} - \mu_{\triangle x}) \tag{3}$$

where x_t^* is the vector of trend outputs and $\mu_{\triangle x} = E(\triangle x_t)$. To obtain a solution for the trend components write the VECM in (2) as

$$\Delta x_t = \mathcal{K}_o H_t + \sum_{j=1}^{k} \theta_j \Delta x_{t-j} + \beta v_{t-1} + \zeta_t. \tag{4}$$

[1] We exclude a linear time trend in the VECM as it would imply a quadratic trend in the levels of the variables.

where $\mathcal{K}_o = (\theta_o, \varkappa)$ and $\varkappa$ contains the κ_{ji} vectors, and:

$$H_t = \begin{bmatrix} \Xi_t \\ \mathcal{D}_t \end{bmatrix}$$

where $\mathcal{D}_t$ contains the $D_{jt-i}s$, and $\upsilon_{t-1} = \alpha' x_{t-1} + \gamma' t \Xi_t$. It follows that:

$$\upsilon_t = \alpha' x_t + \gamma'(t+1)\Xi_t = \alpha' \Delta x_t + \gamma' \Xi_t + \upsilon_{t-1}$$

and then:

$$\upsilon_t = \mathcal{K}_{oo} H_t + \alpha' \theta_1 \Delta x_{t-1} + \cdots + \alpha' \theta_k \Delta x_{t-k} + \left(I + \alpha' \beta\right) \upsilon_{t-1} + \alpha' \zeta_t \qquad (5)$$

where:

$$\mathcal{K}_{oo} = \left(\alpha' \theta_o + \gamma', \alpha' \varkappa\right).$$

Appending (5) to the system in (4) we have a first order stationary vector autoregression of the form:

$$z_t = A_o H_t + A_1 z_{t-1} + \Psi \zeta_t \quad t = 1, \ldots, T \qquad (6)$$

where z_t' is the $(1 \times pk + r)$ vector:

$$z_t' = (\Delta x_t', \Delta x_{t-1}', ..., \Delta x_{t-k+1}', \upsilon_t').$$

The matrices A_o and A_1 are defined as:

$$A_o = \begin{bmatrix} \mathcal{K}_o \\ 0 \\ 0 \\ \vdots \\ 0 \\ \mathcal{K}_{oo} \end{bmatrix}$$

and:

$$A_1 = \begin{bmatrix} \theta_1 & \theta_2 & \cdots & \theta_{k-1} & \theta_k & \beta \\ I & 0 & \cdots & 0 & 0 & 0 \\ 0 & I & \cdots & 0 & 0 & 0 \\ \vdots & \vdots & \cdots & \vdots & \vdots & \vdots \\ 0 & 0 & \cdots & I & 0 & 0 \\ \alpha'\theta_1 & \alpha'\theta_2 & \cdots & \alpha'\theta_{k-1} & \alpha'\theta_k & \alpha'\beta + I \end{bmatrix} \qquad (7)$$

and:

$$\Psi = \begin{bmatrix} I \\ 0 \\ \vdots \\ 0 \\ \alpha' \end{bmatrix}.$$

From (6) it follows that:

$$E(z_t) = \mu_z = (I - A_1)^{-1} A_o H_t \qquad (8)$$

so that:

$$z_t - \mu_z = (I - A_1 L)^{-1} \Psi \zeta_t. \qquad (9)$$

Define the matrix:

$$G = \begin{bmatrix} I_p \\ 0 \\ \vdots \\ 0 \end{bmatrix},$$

then $G'z_t$ selects out Δx_t and it follows from (6) that:

$$\Delta x_t - \mu_{\Delta x} = G'(z_t - \mu_z) = G'(I - A_1 L)^{-1}\Psi\zeta_t = C(L)\zeta_t \tag{10}$$

which is the moving average representation. Inverting $[I - A_1]$, see Attfield [1, (2003)] for fuller details, it is straightforward to show that[2]:

$$C(1) = G'[I - A_1]^{-1}\Psi = \theta(1)^{-1} - \theta(1)^{-1}\beta(\alpha'\theta(1)^{-1}\beta)^{-1}\alpha'\theta(1)^{-1}$$

where $\theta(1) = I_p - \sum_1^k \theta_i$. An alternative well known expression for $C(1)-$ (see, for example, Johansen [12, (p.52, 1995)]) - is:

$$C(1) = \alpha_\perp\left(\beta'_\perp\theta(1)\alpha_\perp\right)^{-1}\beta'_\perp$$

where $\alpha_\perp$ and $\beta_\perp$ are the orthogonal complements of α and β such that $\alpha'_\perp\alpha = 0$ and $\beta'_\perp\beta = 0$.

The expectations term in equation (3), can then be written:

$$\sum_{i=1}^{i=\infty} E_t(\Delta x_{t+i} - \mu_{\Delta x}) = G'A_1[I - A_1]^{-1}(z_t - \mu_z) \tag{11}$$

and some algebra produces:

$$x_t^* = C(1)\theta(L)x_t - Q\gamma'(t+1)\Xi_t + \delta_o \tag{12}$$

where $\theta(L) = I_p - \sum_{j=1}^k \theta_j L^j$ and $Q = \theta(1)^{-1}\beta(\alpha'\theta(1)^{-1}\beta)^{-1}$ and $\delta_o = -C(1)\theta^*(1)\mu_{\Delta x} + Q\mu_v$. The expressions $\mu_{\Delta x}$ and μ_v are the means of stationary variables and can be estimated from sample counterparts.

Equation (12) then contains expressions for permanent stochastic output trends for each of the economies calculated from interactions between them all and allowing for a mean and trend break for each country.

4. Convergence in Output

There are a number of definitions of "convergence" in the time series context. Li and Papell [14, (1999)] classify them as stochastic convergence, deterministic convergence and Bernard-Durlauf convergence. Stochastic convergence requires that the log of relative outputs be trend stationary and coincides with the Ogaki-Park [15, (1997)] definition of stochastic cointegration which implies that the cointegrating vector eliminates stochastic trends in the data series. Deterministic convergence is stronger as the log of relative outputs are required to be stationary and coincides with the Ogaki-Park definition of

[2] Proietti [17, 1997] obtains the same result using the Kalman filter except that instead of $\Theta(1)^{-1}$ he has $(\Theta(1) - \beta\alpha')^{-1}$ but it is easy to show that the two forms give exactly the same $C(1)$.

deterministic cointegration which requires the cointegrating vector to remove both stochastic and deterministic trends. Bernard-Durlauf convergence is strongest of all as the log of relative outputs are required to be stationary with zero mean.

In terms of equation (12) then, the output series, x_t, are given by:

$$x_t = x_t^* + x_t^T \tag{13}$$

where x_t^T is the stationary, mean-zero, transitory component of output. Premultiplying (13) by the set of cointegrating vectors, α', yields:

$$\alpha' x_t^* = -\gamma'(t+1)\Xi_t + \mu_v + \alpha' x_t^T$$

since $\alpha' C(1) = 0$ and $\alpha' Q = I$. So, provided the set of cointegrating vectors satisfy:

$$\alpha' = [-I_{p-1} | \iota_{p-1}]$$

stochastic convergence will always be satisfied. With one break in the time trend (12) can be written:

$$x_t^* = C(1)\theta(L)x_t - Q\gamma_1'(t_1+1) - Q\gamma_2'(t_2+1) + \delta_o$$

and premultiplying by α' :

$$\alpha' x_t^* = -\gamma_1'(t_1+1) - \gamma_2'(t_2+1) + \mu_v$$

and, since we are concerned with whether convergence has occurred during the latter part of the sample we require $\alpha' = [-I_{p-1}|\iota_{p-1}]$ and $\gamma_2' = 0$ for deterministic convergence[3]. Finally, for Bernard-Durlauf convergence we require $\alpha' = [-I_{p-1}|\iota_{p-1}]$, $\gamma_2' = 0$ and that μ_v in the second part of the sample is zero. An expression for μ_v can be obtained as [see Attfield [1, (2003)]]:

$$\mu_v = -(\alpha'\theta(1)^{-1}\beta)^{-1} \left[(\alpha'\theta(1)^{-1}\theta_o + \gamma')\Xi_t + \alpha'\theta(1)^{-1}\varkappa\mathcal{D}_t \right]$$

so that we would require $\gamma_2' = 0$ and θ_{o2}, the intercepts in the VECM for the second part of the sample, to be equal to 0.[4]

Even if the output series have not converged they may still be generated by a single, common stochastic trend. That is, for p countries if there are $(p-1)$ cointegrating equations we can apply normalisation conditions to write, without any loss of generality:

$$\alpha' = [-I_{p-1}|a]$$

where I_{p-1} is the identity matrix of order $p-1$ and a is a $(p-1 \times 1)$ vector of cointegrating coefficients. With this specification each of $(p-1)$ country's output is a function of one country, the pth. In this general case $\alpha_\perp' = (a', 1)$ and it follows that:

$$\begin{aligned} C(1)\theta(L)x_t &= \alpha_\perp \left(\beta_\perp' \Theta(1)\alpha_\perp\right)^{-1} \beta_\perp' \theta(L) x_t \\ &= \begin{bmatrix} a \\ 1 \end{bmatrix} f_t \end{aligned}$$

[3]This definition of deterministic convergence doesn't quite coincide with the Ogaki-Park definition of deterministic cointegration as it isn't the cointegrating vector which eliminates the deterministic trend. Nevertheless, with the restrictions on the vector of cointegrating coefficients and the restriction that $\gamma_2' = 0$, in the latter part of the sample, the log differences have neither stochastic nor deterministic trends.

[4]We ignore restrictions on the parameters in $\varkappa$ as they are attached to variables which have only been included in the model to eliminate observations at the start of each of the periods after a break point.

where $f_t = (\beta_\perp' \Theta(1)\alpha_\perp)^{-1} \beta_\perp' \theta(L) x_t$ is the single stochastic trend and equation (12) becomes:

$$x_t^* = \begin{bmatrix} a \\ 1 \end{bmatrix} f_t - Q\gamma'(t+1)\Xi_t + \delta_o.$$

If a is restricted to a vector of 1's we revert to the cases discussed above. Although we have restricted the discussion to the trend derived from the Beveridge-Nelson decomposition the results would carry over to the Gonzalo-Granger [11, (1995)] decomposition where, omitting constants and trends, the permanent component is given by $\alpha_\perp(\beta_\perp'\alpha_\perp)^{-1}\beta_\perp' x_t$ and Proietti [17, (1997)] where the permanent component is $C(1)\theta(1)x_t$. In both cases the permanent component is annihilated by pre-multiplication by α' so that stochastic convergence is satisfied with $\alpha' = [-I_{p-1}|\iota_{p-1}]$. Deterministic and Bernard-Durlauf convergence would require zero trend coefficients and constants as discussed above.

5. Empirical Results

5.1. Data Series. In this section we model quarterly, seasonally adjusted real output, Y_t, for the EU members Belgium, Finland, France, Holland, Italy, Spain and the UK for the period 1980Q1 to 2001Q3. The series were obtained from the OECD Main Economic indicators data bank at MIMAS (Manchester Information Service at Manchester University) using the Timeweb Explorer interface. Output is defined as gross domestic product minus government final consumption expenditure. All series are quarterly annualised seasonally adjusted and in constant, 1995, prices. For comparison across members of the EU, all values were converted to US dollars using the $/£ exchange rate end period. Annual population data were obtained from the U.S. Bureau of the Census, International Data Base and log-linearly interpolated to quarterly series[5]. The EU countries excluded from the analysis are Greece, Ireland, Luxembourg, Austria, Denmark, Germany, Portugal and Sweden because data on the series are not available prior to 1988Q1. The starting point of 1980Q1 was chosen as this included Belgium and Spain where series are not available until that date.

5.2. Results. For each country the logarithm of output was tested for a unit root and although the null of a unit root could not be rejected for any of the variables using standard ADF tests, to allow for structural breaks which may affect the standard tests, the series were tested using the procedure of Banerjee et al [3, (1992, Table2, p.278)] which allows for a jump or break in the intercept, a mean shift, or a change in the slope of the trend, a trend shift. The null of a unit root could not be rejected[6]. It follows that x_t, the vector of all log outputs is $I(1)$ and we can next test for cointegration.

For the hypothesis of a single stochastic trend driving output for all 7 countries we require 6 cointegrating vectors. For the null hypothesis of deterministic convergence we require $\alpha' = [-I_6|\iota_6]$ and $\gamma_2' = 0$. Using the standard Johansen maximum likelihood test, [12, (1995)], for cointegrating rank, without assuming any structural breaks, we found we could only accept 4 cointegrating equations at the 5% level or 5 at the 10%

[5] Neither per-capita nor PPP adjusted quarterly series are available from the OECD source.

[6] Test statistics allowing for mean shifts were for Belgium, Finland, France, Italy, Netherlands, Spain and UK respectively: –4.19, 3.45, -4.26, -4.64, -4.20, -4.22, -4.16 using the BIC to select lag length. The 5% critical value is -4.8. For trend shifts test statistics were -4.04, -2.88, -3.80, -3.79, -3.91, -3.57, -3.30. The critical value for a shift in trend is -4.48. Critical values were obtained from Banerjee et al [3, (1992, Table2, p.278)]. Results were obtained using the GAUSS programming language, [9, 2001].

level in a model with unrestricted constants [7]. The reason for this result could be that structural breaks in the series, which make the cointegrating equations trend-stationary, influence the test statistics in favour of rejecting the stationarity hypothesis[8].

To use the formulation in (2) to test for cointegration when there are structural breaks, we first have to identify the break points in the system. There are a number of papers which suggest methods for finding break points in single equation cointegrating models (e.g., Gregory and Hansen [10, (1996)], Bai and Perron [6, (1998)]) while BLS, [2, (1998)] give a method for estimating confidence intervals for break dates in multivariate systems and argue that tighter intervals can be obtained from a multivariate approach. Their method assumes the system in (1) with given cointegrating vectors and estimates a confidence interval for a shift in the intercept in the VECM. Although this isn't exactly the same as the model in (2), because there are no trends in the cointegrating equations, it provides a starting point for finding breaks. Using the prior $\alpha' = [-I_6|\iota_6]$ the BLS procedure located a break point at 1992Q4 while using MLE estimates (without trend breaks) it located 1993Q3 with a 10% confidence interval of (1993Q2,1993Q4)[9]. We assumed the break point of 1993Q3, although the model is fairly robust to dates around this date.

Having identified the break points we can now test for cointegration as Johansen et al [13, (2000)] derive the distribution of the trace test statistic for testing for the rank of the cointegrating space in a model such as equation (2). They also calculate the weights for the estimated response surface to enable critical values to be easily calculated from a $\Gamma-$ distribution. Applying these methods[10] to the model in (2) we obtained the results in Table 1.

Table 1. Tests of Rank

Hypothesis	Test Statistic	p-value
$r \leq 4$	60.358	0.033
$r \leq 5$	35.110	0.077
$r \leq 6$	14.581	0.176

Allowing for the break point the likelihood rank test statistic rejects 5 cointegrating vector in favour of 6 at around 8% so that we can conclude that the rank of the cointegrating space is 6 and, without any loss of generality, normalise the set of cointegrating vectors as:

$$\alpha' = [-I_{p-1}|a]$$

[7]With one lag first difference in the VECM the "trace" test statistic for $Ho{:}r \leq 3, H_A : r \leq 4$ was 49.48 with critical values, from Osterwald-Lenum [16, 1992], of 43.95 (10%) and 47.21 (5%). For $Ho{:}r \leq 4, H_A : r \leq 5$ the test statistic was 28.9 with critical values 26.79 (10%) and 29.68 (5%). For the case of a trend included in the cointegrating equations the results were:
$Ho{:}r \leq 3, H_A : r \leq 4$ 64.86 (59.14, 62.99)
$Ho{:}r \leq 4, H_A : r \leq 5$ 41.35 (39.06, 42.44).

[8]Neither the AIC, BIC or LR testing down procedures were very useful for selecting the lag length because of the very flat likelihood surface. Moreover, in the likelihood procedure with broken trends, it was found that the companion matrix for the VECM had roots outside the unit circle with lags of first differences of two or more. In all the multivariate procedures in the paper, therefore, we employed only one lag of first differences.

[9]The lExp-W and Sup-W test statistics were 62.4 (35.1) and 27.1 (12.2), for the imposed cointegrating vectors of $\alpha' = [-I_6|\iota_6]$ and 38.4 (35.1) and 15.3 (12.2). The 5% critical values for the statistics (given in parenthesis) were obtained by simulations similar to those reported in BLS.

[10]All the results in this paper were obtained using the GAUSS[9, (2001)] programming language. For the multivariate cases BIC and AIC were consistent in selecting one lag of first differences in the VECM - two lags in levels.

as there is only one stochastic trend driving the output of the seven economies. Table 2(a) gives the results for the unrestricted model. The UK was chosen as numeraire so that each cointegrating equation relates the log of output for each country to the log of output for the UK. The coefficient between the log of each country's output and UK output, the elements of the vector a, are not significantly different from unity for Belgium, Finland, France, the Netherlands and UK but are significantly greater than unity for Italy and Spain for any conventional significance level. All time trend coefficients for all equations and both trend periods are significantly greater than zero. Table 2(b) gives the estimates for the test of stochastic convergence with the restrictions $a = \iota$, which is rejected at both 5% and 1% with a LR test with a p-value of 0.005[11]. This is not surprising given the individual results for Italy and Spain. With these two countries coefficients unrestricted the LR test statistic is 13.33, with 4 degrees of freedom, which has a p-value of 0.01 so that stochastic convergence for the five countries, Belgium, Finland, France, the Netherlands and the UK, is on the borderline for acceptance. Table 2(c) gives the results of the test for deterministic convergence where we require $a = \iota$ and that all the coefficients on the trend in the second period are zero. The LR test statistic is 38.82 which rejects deterministic convergence at any conventional level. Given the rejection of deterministic convergence we did not test for the even stronger Bernard-Durlauf convergence which imposes the additional restrictions that the intercepts in the VECM are zero in the second period of the sample.

Figures 1 to 7 plot the stochastic trend for each country against its actual output using the model in Table 2(b) with $a = \iota$ so that their long run stochastic trends only differ because of the coefficients on the broken time-trends and intercept terms i.e., the case of stochastic convergence. Figure 8 plots the long run stochastic trend for each country for the second break period, i.e., 1994 to 2001 and gives a visual summary of the results with the UK, Belgium, Finland, the Netherlands and France close to convergence, and only really distinct because of the effect of the magnitude of the coefficients on the time trends.

6. Conclusion

Using quarterly, OECD output series from 1980Q1 to 2001Q3 for seven European economies, Belgium, Finland, France, Italy, the Netherlands, Spain and the UK, this paper finds some evidence for stochastic convergence in output growth in that Belgium, Finland, France, the Netherlands and the UK have identical long run stochastic trends when we allow for a shift in the mean of the VECM and in the time trend in the cointegrating equations in 1993Q3. If we assume time trends in the cointegrating equations are transitory, so that in the long run their coefficients are zero, we could accept strong deterministic convergence for this same group of five countries. If time trends are considered permanent then we can reject the hypothesis of strong convergence. The findings on convergence are not wholly consistent with those of Bernard and Durlauf [4, (1995)] in their study of 15 countries where they found evidence for "3 to 6" common trends driving the output series. Their data series finish in 1987 and are for a wider cross section of countries than those studied here and do not incorporate structural breaks within the cointegrating equations. The results are however, consistent with Strazicich and Lee [18, (2001)] and Li and Papell [14, (1999)] who do find evidence for weak convergence for annual data series up to the mid 90's when they incorporate endogenous breaks in their tests for cointegration.

[11] The restricted model and asymptotic standard errors were estimated using the switching algorithm technique of Doornik [8, (1995)].

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Tests of European Output Series

Table 2(a). Unrestricted model

	Belgium	Finland	France	Italy	Netherlands	Spain	UK
Real $ output per capita as a proportion of output of UK (a_j)	1.201	1.206	1.182	1.224	0.973	1.487	1
(Estimated standard error)	(0.115)	(0.130)	(0.115)	(0.072)	(0.109)	(0.104)	
Coefficient on trend period 1	0.711	-0.658	0.300	-0.300	1.086	-0.348	-
(Estimated standard error)	(0.132)	(0.148)	(0.115)	(0.082)	(0.124)	(0.119)	
Coefficient on trend period 2	-1.229	-0.988	-1.141	-1.194	-0.960	-1.359	-
(Estimated standard error)	(0.265)	(0.300)	(0.265)	(0.166)	(0.250)	(0.240)	

Table 2(b). Ratios Restricted to Unity

	Belgium	Finland	France	Italy	Netherlands	Spain	UK
Real $ output per capita as a proportion of output of UK (a_j)	1	1	1	1	1	1	1
Coefficient on trend period 1	0.731	-0.643	0.317	-0.288	1.097	-0.314	-
(Estimated standard error)	(0.121)	(0.141)	(0.119)	(0.085)	(0.114)	(0.145)	
Coefficient on trend period 2	-1.261	-1.016	-1.174	-1.221	-0.969	-1.422	-
(Estimated standard error)	(0.263)	(0.308)	(0.260)	(0.186)	(0.247)	(0.315)	

Likelihood ratio (LR) test of:	Ho: $a_j = 1$ for all j
LR test statistic, $\chi^2(6)$:	18.79
(p-val)	(0.005)

Table 2(c). Ratios restricted to Unity and Trend in Period 2 Restricted to Zero

	Belgium	Finland	France	Italy	Netherlands	Spain	UK
Real $ output per capita as a proportion of output of UK (a_j)	1	1	1	1	1	1	1
Coefficient on trend period 1	0.742	-0.621	0.326	-0.270	1.105	-0.298	-
(Estimated standard error)	(0.275)	(0.208)	(0.262)	(0.239)	(0.226)	(0.299)	
Coefficient on trend period 2	0	0	0	0	0	0	-

Likelihood ratio (LR) test of:	Ho: $a_j = 1$ for all j and period 2 trend coefficients restricted to zero
LR test statistic, $\chi^2(12)$:	38.82
(p-val)	(0.0001)

Fig.1 Comparison of Trends in Real Per Capita Output for Belgium

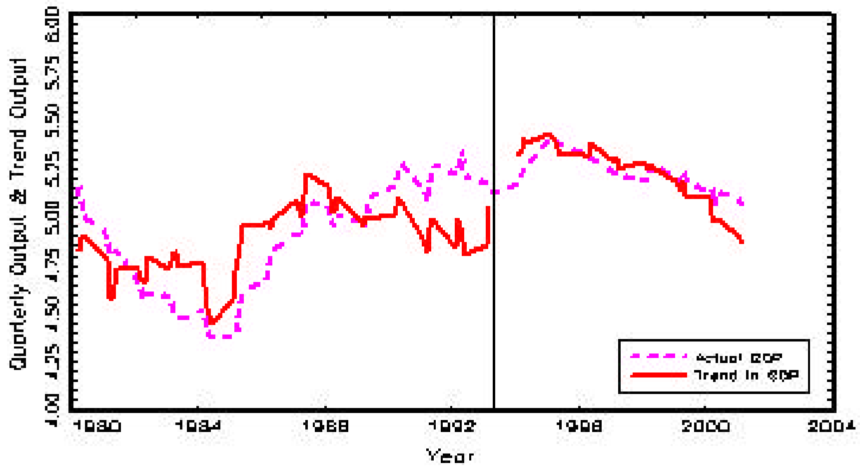

Fig.2 Comparison of Trends in Real Per Capita Output for Finland

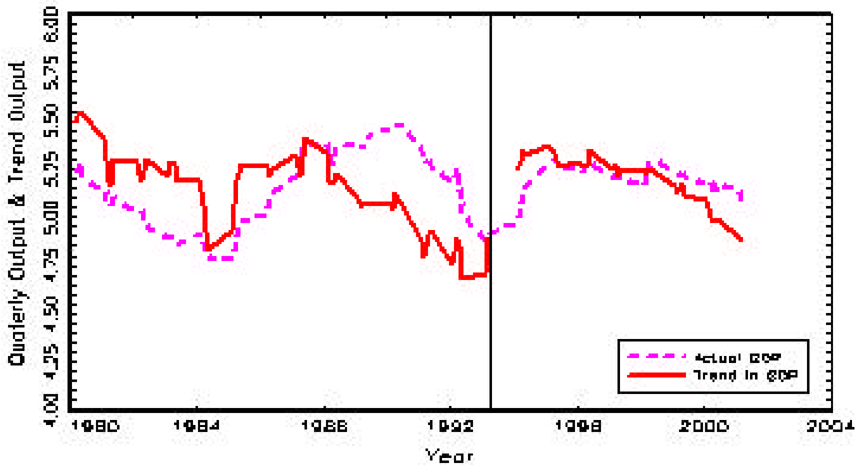

Fig.3 Comparison of Trends in Real Per Capita Output for France

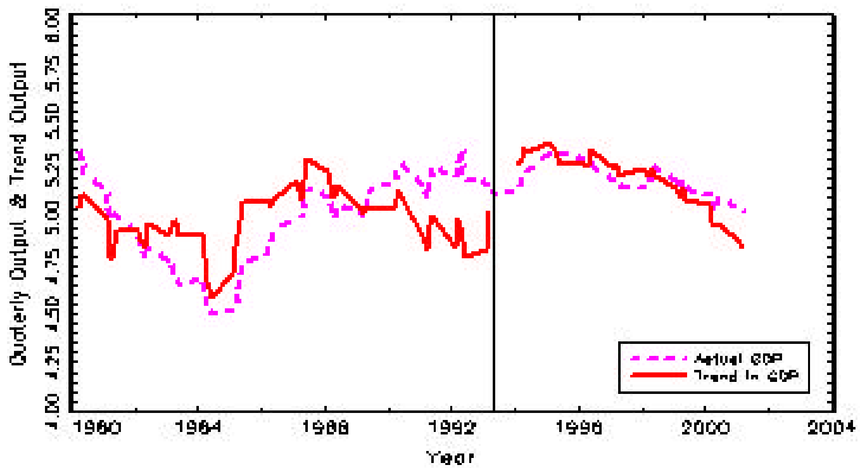

Fig.4 Comparison of Trends in Real Per Capita Output for Italy

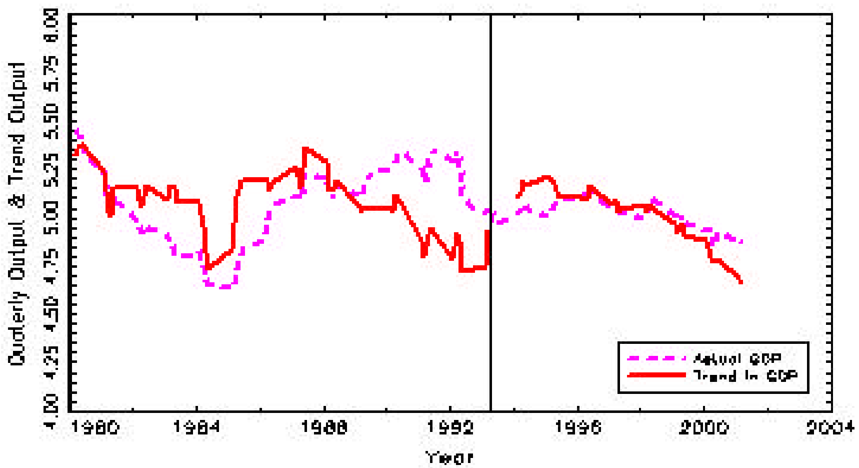

Fig.5 Comparison of Trends in Real Per Capita Output for Netherlands

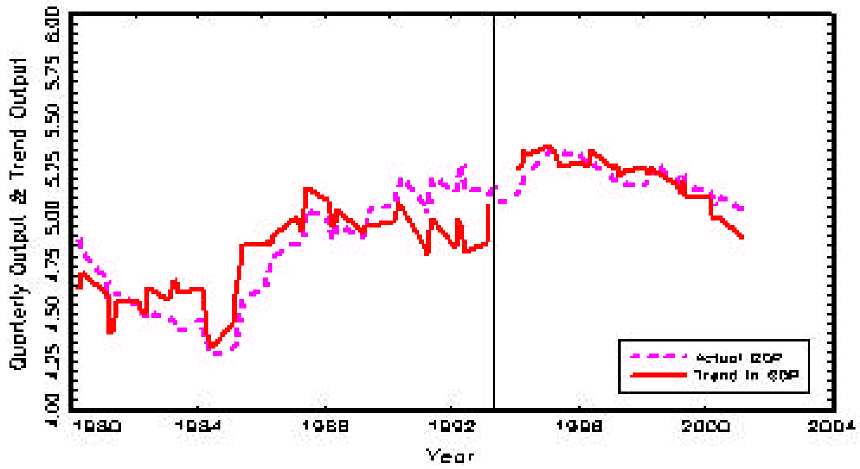

Fig.6 Comparison of Trends in Real Per Capita Output for Spain

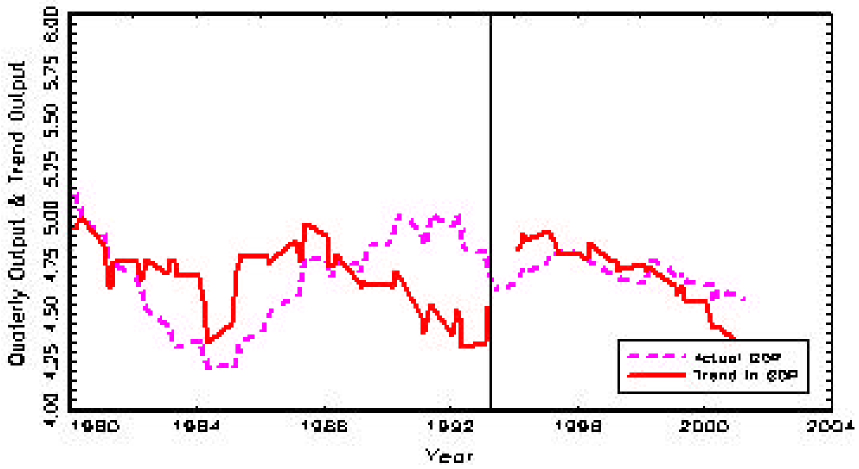

Fig.7 Comparison of Trends in Real Per Capita Output for UK

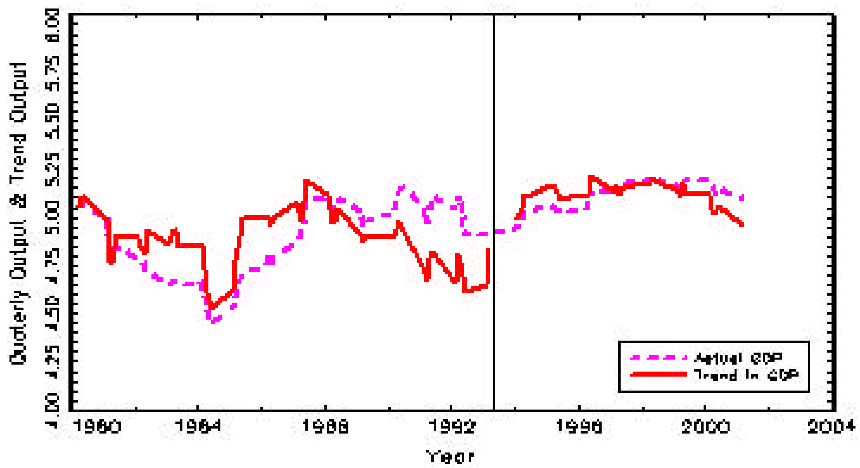

Fig.8 Comparison of Trends in Real Per Capita Output for All Countries