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Nonlinear Dynamics, Spillovers and Growth in the G7 Economies: An Empirical Investigation

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This paper proposes an empirical growth model which is consistent with a stochastic steady-state labour productivity level varying over time and across countries, where the disequilibrium mechanism leading to long-run equilibrium follows a nonlinear equilibrium correction model. Using data for the G7 economies during the postwar period since 1950, the empirical analysis yields a long-run model which implies plausible estimates of the production function parameters. Postwar economic growth in each of the G7 countries appears to be well characterized by a nonlinear equilibrium correction model where the dynamic adjustment towards long-run equilibrium is governed by a logistic function, while also capturing spillover effects in growth dynamics.

INTRODUCTION

This paper examines empirically the determinants of both the steady state and the dynamic adjustment path towards the steady state in the context of a variant of the augmented Solow–Swan (Solow 1956; Swan 1956; Mankiw *et al.* 1992) model using postwar data for the G7 economies. Several recent papers have highlighted the fact that many of the early studies investigating growth dynamics using cross-sectional regressions are subject to serious econometric difficulties, directing researchers towards econometric techniques that allow cross-country heterogeneity.[1] This paper employs time-series methods which allow us to accommodate the presence of cross-country heterogeneity both in the steady state and in the dynamic adjustment path. The empirical model proposed here is derived from extending the traditional augmented Solow–Swan model by allowing for a stochastic element, with the steady-state equilibrium level of labour productivity modelled as a stochastic process, varying both over time and across countries, determined by the level of technology, the rate of capital accumulation and the rate of labour force growth.

Although some recent research recognizes the importance of allowing the determinants of the steady state to be stochastic, there is —to the best of my knowledge—no empirical study in the relevant literature that tests the augmented Solow–Swan model with each of the determinants of the steady state allowed to be stochastic. In other words, the assumption that at least one of these—the technology level, the rate of capital accumulation, the rate of labour force growth—is non-stochastic is routinely made in the empirical literature, and this may, at least partially, explain the failure to detect cointegration in the static regression implied by the augmented Solow–Swan model (e.g. Cellini 1997; Sarno 1999).

Further, the augmented Solow–Swan model predicts that labour productivity is governed by a law of motion which suggests nonlinear adjustment towards long-run equilibrium; this is, however, a theoretical implication largely neglected by the relevant empirical literature to date. This theoretical prediction appears to be consistent with the strong empirical evidence reported by the recent literature (discussed below), suggesting that the dynamics of postwar business cycles in the G7 countries display significant nonlinearities. Thus, one might expect that allowing for nonlinear adjustment towards equilibrium in the empirical modelling of economic growth may yield some improvement upon the corresponding linear equations.

The present paper contributes to the relevant literature in that I propose an empirical growth model which may be viewed as a generalization of the augmented Solow–Swan model, explicitly allowing for nonlinear adjustment towards the long-run equilibrium as well as for spillovers in growth dynamics across countries. Using postwar data for the G7 countries since 1950, my analysis ultimately results in the estimation of a nonlinear equilibrium correction model (ECM), where each of the variables determining the steady state is treated as stochastic, the cointegrating vector implies plausible estimates of the production function parameters, and the adjustment towards equilibrium is found to be particularly well characterized by a logistic cumulative distribution function of the type proposed by Granger and Swanson (1996). The empirical model allows for the dynamics of labour productivity in one country to be affected by the equilibrium errors from other countries, suggesting the existence of significant spillover effects in growth dynamics within the G7.

The remainder of the paper is set out as follows. Section I outlines the theoretical growth literature motivating my empirical work, in addition to discussing some relevant empirical literature. Section II describes my empirical modelling strategy, briefly introducing the recently developed theory of equilibrium correction models for nonlinear attractors. Section III describes the data. Section IV reports and discusses the empirical results. A final section concludes.

I. Modelling Economic Growth

In this section I briefly make reference to the theoretical growth literature that is relevant to this paper and put the specifications employed in the empirical analysis discussed below in context.

Consider a constant-returns-to-scale production function, for which the Inada conditions hold:

$$(1) \qquad Y_t = K_t^{\alpha} H_t^{\beta} (A_t L_t)^{\gamma},$$

where Y, K, H and L denote output, the level of physical capital, the level of human capital and labour, respectively; A is a labour-augmenting technological shift parameter; α and β are technology parameters, $\gamma \equiv 1 - \alpha - \beta$; t is a time subscript. Under standard assumptions on physical and human capital accumulation (both depreciating at a rate δ), and assuming that the labour force grows at a rate n and technological progress at a rate g, one can derive the constant steady-state level for K/AL, H/AL and Y/AL:

$$(2)\qquad \left(\frac{K}{AL}\right)^* = \hat{k}^* = \left[\frac{s_K^{1-\beta} s_H^{\beta}}{\zeta}\right]$$

$$\left(\frac{H}{AL}\right)^* = \hat{h}^* = \left[\frac{s_K^{\alpha} s_H^{1-\alpha}}{\zeta}\right]$$

$$\left(\frac{Y}{AL}\right)^* = \hat{y}^* = \left[\frac{s_K^{\alpha/\gamma} s_H^{\beta/\gamma}}{\zeta^{(\alpha+\beta)/\gamma}}\right],$$

where s_K and s_H denote the fractions of output devoted to the accumulation of physical and human capital respectively; $\zeta \equiv n + g + \delta$; and carets and asterisks refer to variables defined in efficiency units and to steady-state variables, respectively. The implied steady-state productivity equation is then of the form

$$(3)\qquad \ln y_t^* = \ln A + (\alpha/\beta)\ln s_K + (\beta/\gamma)\ln s_H - [(\alpha+\beta)/\gamma]\ln \zeta,$$

with $\ln A \equiv \ln A_0 + gt$ (see e.g. Mankiw *et al.* 1992 and the references therein).

While a large literature has examined the Solow–Swan model or variants of it in a deterministic framework (Temple 1999), a recent literature has started to consider reformulations of the Solow–Swan model which allow for stochastic—as opposed to constant—level of technology, rate of capital accumulation and rate of labour force growth. Notably, Lee *et al.* (1997, 1998) set out formally the implications of a stochastic Solow–Swan model where the processes describing technology and employment are stochastic (although savings rates are fixed), showing that the model's implications become rather different from its conventional deterministic version. Lee *et al.* demonstrate, *inter alia*, that traditional estimates of 'beta convergence' may be subject to substantial biases.[2]

Although the assumption of homogeneity of g across the countries of the OECD (hence the G7) is tenable in the light of the recent literature, suggesting that the rate of advancement of knowledge is not country-specific across major industrialized countries,[3] the assumption of non-stochastic $\ln A$, s_K, s_H and n is unnecessary. If $\ln A$, s_K, s_H and n vary over time and across countries, the augmented Solow–Swan model implies that steady-state labour productivity also varies over time and across countries, and in the neighbourhood of the steady-state path, labour productivity of country j is governed by the equation

$$(4)\qquad \begin{aligned}\ln y_{jt+1} - \ln y_{jt} &= g + (1 - e^{-\theta_{jt}})\{(\ln A_{j0} + gt) + (\alpha/\gamma)\ln s_{K_{jt}} \\ &\quad + (\beta/\gamma)\ln s_{H_{jt}} - [(\alpha+\beta)/\gamma]\ln \zeta_{jt} - \ln y_{jt}\} \\ &= g + (1 - e^{-\theta_{jt}})\{\ln y^* - \ln y\}_{jt},\end{aligned}$$

where $\theta_{jt} \equiv \zeta_{jt}\gamma$ and $\zeta_{jt} \equiv (n_{jt} + g + \delta)$ (see Mankiw *et al.* 1992; Durlauf and Johnson 1995; Cellini 1997). It is clear that, unlike the model proposed by Lee *et al.* (1997, 1998), (4) is effectively derived by adding a stochastic element to the basic setup of the deterministic Solow–Swan model without building stochastic variation from the outset. Nevertheless, (4) provides interesting insights and has a natural interpretation for empirical growth modelling. In particular, it may be viewed as a nonlinear ECM such that labour productivity in a country rises (falls) if its current level is below (above) its steady-state level with a time-varying speed of adjustment towards the steady state.

Although the Solow–Swan growth model is derived within a closed-economy context and hence neglects international borrowing and lending and international trade—while being less sophisticated than the Ramsey–Cass–Koopmans closed-economy model of joint determination of growth and savings rates (Ramsey 1928; Cass 1965; Koopmans 1965)—it still conveys important insights on economic growth. Nevertheless, in order to reconcile growth models with the data, it may be important to consider global linkages across countries and international business or growth cycles, given the stylized fact that business cycles are highly correlated across developed countries. (See Baxter, 1995, for a survey of the main findings in this context.) Technology spillovers seem to be the most important determinant of these correlations, and they are strong enough that the opportunities for international risk-sharing are rather limited (see Baxter 1995)—related studies are Durlauf (1993), Coe and Helpman (1995), Quah (1996b), Keller (1998), Bayoumi *et al.* (1999a,b). In an interesting related paper, Lee (1998) investigates cross-country growth interdependencies within the G7, motivating his empirical analysis on the basis of a model where interdependencies are generated by the equilibrating process initiated when countries experience current account disequilibria.

While these models may be used to motivate relationships between output growth dynamics as well as between output levels across countries, in this paper I focus on the effects of spillovers in growth dynamics. Defining for simplicity the nonlinear function governing the adjustment towards equilibrium $F(\cdot)$ and assuming an N-country world, consider the following generalization of the nonlinear ECM governing the labour productivity dynamics of country $j \in \{1, 2, ..., N\}$:

$$\ln y_{jt+1} - \ln y_{jt} = g + F[v_1(\ln y^* - \ln y)_{1t}, v_2(\ln y^* - \ln y)_{2t}, ..., v_N(\ln y^* - \ln y)_{Nt}], \tag{5}$$

where stability (convergence to the steady state) requires a non-zero response to $(\ln y^* - \ln y)_{jt}$ (i.e. $v_j \neq 0$), while a non-zero response to $(\ln y^* - \ln y)_{it}$ $(i = 1, 2, ..., N;\ i \neq j)$ implies the existence of spillover effects between country j and country i. Although the allowance for spillover effects in (5) is rationalized on the basis of stylized facts and our economic prior of the existence of significant spillover effects rather than directly derived from the Solow–Swan theory underlying this work, the principal motivation of this paper is empirical and I consider this allowance an important element of the empirical modelling of economic growth. The absence of short-run relationships between one country's (say j's) labour productivity law of motion and deviations from the steady state in other countries would show up in the lack of statistical significance of the speed of adjustment parameters associated with $(\ln y^* - \ln y)_{jt}$ in $F(\cdot)$ for $i \neq j$ and, therefore, the collapse of equation (5) to the law of motion (4).

Given (3)–(5), equilibrium correction and cointegration techniques seem to be a natural way of testing the implications of the empirical model presented here. Establishing cointegration between labour productivity and its determinants implies the stationarity of the difference between the steady-state level of labour productivity and its current level as well as the stationarity of labour productivity growth, also suggesting the existence of an ECM (Granger 1986; Engle and Granger 1987). Granger and Swanson (1996) strongly emphasize the importance of investigating the (linear or nonlinear) ECM implied by a cointegrating relationship since cointegration is 'just a property' (p. 543), whereas an ECM is a potential data-generating mechanism.

As mentioned earlier, previous empirical studies have typically kept fixed at least one of the determinants of the steady state implied by the augmented Solow–Swan model, generating a potential omitted-variables problem. Also, a further possible rationalization of the difficulty in establishing cointegration in this context is that the span of available data may be too short to provide sufficient test power in conventional unit root tests (Shiller and Perron 1985). This may be particularly important if the adjustment towards equilibrium is slow—e.g. if shocks to the steady-state level of labour productivity are highly persistent—and if the true adjustment towards equilibrium occurs nonlinearly, which may be expected given equations (4)–(5) (see Balke and Fomby 1997; Enders and Granger 1998). Granger and Swanson (1996), for example, update the data set used by King *et al.* (1991)—who found cointegration between labour productivity and total factor productivity using those data—and detect strong nonlinearity in the dynamic relationships between the time series examined.[4] Kontolemis (1997) also provides strong evidence of asymmetry of business cycles in the G7 countries. Sarno (1999) detects evidence of nonlinearity in postwar labour productivity growth of the G7 countries, which is captured using a logistic smooth transition autoregressive model, and a natural extension of this work is to employ nonlinear equilibrium correction models.

The present paper may be seen as an extension of this strand of the literature, in that here I estimate the stochastic steady state for each of the G7 countries using postwar data since 1950, and then estimate nonlinear equilibrium correction models with the adjustment towards long-run equilibrium modelled by adopting economically interpretable parametric specifications, while also allowing for spillover effects in growth dynamics within the G7.

II. Nonlinear Econometric Techniques

The starting point of the modelling strategy adopted in this paper is the generalization of linear cointegration and equilibrium correction in a nonlinear setting.[5] Following Granger and Swanson (1996), one may consider a q-component vector of $I(1)$ variables $\mathbf{W}_t = (W_{1t}\ W_{2t}\ \ldots\ W_{qt})'$ with $\mathbf{V}_t = \boldsymbol{\kappa}_t'\mathbf{W}_t$ a r-component vector of $I(0)$ zero-mean processes. A common interpretation is the case when $\boldsymbol{\kappa}_t'\mathbf{W}_t = 0$ is the attractor (equilibrium) of the system, and $||\mathbf{V}_t||$ may be seen as a measure of the extent to which the system is out of equilibrium. Assuming $\mathbf{h}(\cdot) = \mathbf{h}(\boldsymbol{\lambda}'\mathbf{V}_t)$, a nonlinear ECM may then be derived as follows:

$$(6) \qquad \Delta\mathbf{W}_{kt} = \sum_{j=1}^{r} \mu_{kj}\mathbf{h}_j(\boldsymbol{\lambda}_j'\mathbf{V}_{t-1}) + \omega_{\kappa t}, \qquad k = 1, 2, \ldots, q,$$

where Δ denotes the first difference operator, $\omega_{\kappa t}$ is white noise, and $\mathbf{h}(\cdot)$ is a function such that $\mathbf{h}(0) = 0$ and $E[\mathbf{h}(\mathbf{V}_t)]$ exists—note that (6) is obviously balanced, since all terms are short memory in mean processes. Using obvious matrix notation, (6) may be rewritten as

$$(7) \qquad \Delta\mathbf{W}_t = \boldsymbol{\mu}\mathbf{h}(\boldsymbol{\lambda}'\mathbf{V}_{t-1}) + \omega_t,$$

with q equations, r factors on the right-hand side and, therefore, $(q-r)$ independent linear combinations $\mathbf{b}_t = \boldsymbol{\mu} \perp \mathbf{W}_t$ such that $\Delta\mathbf{b}_t = \boldsymbol{\mu} \perp \boldsymbol{\omega}_t$ and each element of $\mathbf{W}_t$ is a random walk.

An interesting extension of cointegration and equilibrium correction modelling involves the 'amalgamation' of cointegrating relationships across different sectors of an economy. For example, let $\mathbf{X}_{1t}$ and $\mathbf{X}_{2t}$ be two q-component vectors of economic variables from different sectors of the economy that cointegrate individually with $\mathbf{V}_{1t} = \boldsymbol{\kappa}_1'\mathbf{X}_{1t}$ and $\mathbf{V}_{2t} = \boldsymbol{\kappa}_2'\mathbf{X}_{2t}$, respectively. If the two sectors of the economy analysed are 'separated in the long run', theoretically $\mathbf{V}_{1t}$ and $\mathbf{V}_{2t}$ are all that would be found, even if the variables from the two different sectors were analysed as a full system. Even if there is long-run separation, however, there may be important short-run relationships between the two sets of variables, and therefore the equilibrium error from one sector may enter the (linear or nonlinear) equilibrium correction equation of the other sector (Konishi and Granger 1992; Granger and Swanson 1996; Granger and Haldrup 1997). This amalgamation may be generalized to the case of cointegration analysis across different regions of a country, or different countries of the world economy. In the context of this study, for example, economic growth may be separated in the long-run across the G7 countries and cointegration may be established in a plausible static cointegrating equation for each of the individual countries. Despite long-run separation, however, the individual short-run relationships may be characterized by the equilibrium error from one equation (country) entering another equilibrium correction equation (country) of the system (G7 economies). This is the approach followed in this paper, where I start by estimating cointegrating relationships and, therefore, equilibrium correction terms, which imply plausible production function parameters. Thus, I estimate a nonlinear ECM for each country of the G7, allowing the deviation from equilibrium in other G7 countries to enter each equilibrium correction equation in order to investigate the possibility of spillover effects, and experiment with various nonlinear functions for $\mathbf{h}(\cdot)$.

III. DATA

Annual data for real gross domestic product (GDP) per worker at a constant international price level, real GDP per capita at a constant international price level, investment share of GDP and population were taken from the Summers–Heston (1991) Penn World Tables.[6] In addition, annual data for the secondary school enrolment rate and population for different age groups were taken from the *Statistical Yearbook* of the UN Educational, Scientific and Cultural Organization (UNESCO). Also, data for the total expenditure share of GDP for research and development (R&D) were taken from the *Basic Science and Technology Statistics* CD of the Organisation for Economic Cooperation and Development (OECD). All data cover the sample period 1950–92 for each of the G7 countries. Using these data, I constructed the data-set employed in the empirical analysis: the natural logarithm of real GDP per worker at constant international price (*Y*); the natural logarithm of the stock of cumulated past total R&D expenditure shares of GDP based on a 15% per annum depreciation assumption (*A*) (see Griliches 1995, 1998);[7] the natural logarithm of the investment share of GDP divided by 100 (*SK*); the natural logarithm of the secondary school enrolment for the appropriate age group divided by 1,000, used as a proxy of the propensity to human capital accumulation (*SH*) (see Mankiw *et al.* 1992, p. 419; Barro and Lee 1993); and the natural logarithm of the sum of the employment growth rate plus the technological progress rate plus the depreciation rate (*Z*).[8]

IV. Empirical Analysis

(a) Unit root tests and cointegration analysis

The empirical results from executing several unit root test statistics on *Y*, *A*, *SK*, *SH* and *Z* (not reported, to conserve space) did not enable me to reject the null hypothesis of non-stationarity for any of the series in levels—with the exception of *SK* for the United States—while I was able to reject the null of non-stationarity for each of the series in first difference. So, I concluded that the series in question are realizations from stochastic processes integrated of order one, $I(1)$.

In order to establish a series for the deviations from the long-run equilibrium of labour productivity, for each of the countries examined, I tested for cointegration in the static regression:

$$Y_t = \phi_0 + \phi_1 t + \phi_2 A_t + \phi_3 SK_t + \phi_4 SH_t + \phi_5 Z_t + u_t. \tag{8}$$

The Monte Carlo results provided by Balke and Fomby (1997) suggest that estimation of the long-run equilibrium using the Johansen (1988, 1991) cointegration procedure when the true adjustment towards equilibrium is nonlinear does not yield misleading results in terms of significant loss of power or size distortion, and therefore I employ this procedure to test for cointegration and to estimate the equilibrium error. Table 1 reports the estimated eigenvalues (panel (a)), the maximum eigenvalue statistics (panel (b)) and the trace statistics (panel (c)) from employing the Johansen procedure in a second-order vector autoregression (VAR) comprising *Y*, *A*, *SK*, *SH*, *Z*, an unrestricted constant term and a deterministic trend restricted to the cointegration space. The results—with and without using Reimers's (1992) adjustment for degrees of freedom—clearly suggest that, for each of the countries examined, there is a long-run relationship between the variables considered, and the cointegrating vector appears to be unique at the 5% significance level.

I then re-executed the Johansen cointegration analysis under the restriction that the rank of the long-run matrix $\mathbf{qb}' = \Pi$ equals unity and that A_t, SK_t, SH_t and Z_t are weakly exogenous for the cointegrating vector $\mathbf{b}$, leaving the cointegrating vector unrestricted (other than normalization on Y_t to -1). This implies that the speed of adjustment coefficients in the first (and only) column of $\mathbf{q}$ associated with A_t, SK_t, SH_t and Z_t are restricted to zero, so that the cointegrating relationship enters only the equation for *Y*; i.e., $\mathbf{q} = [q_Y, 0, 0, 0, 0]'$. Panel (d) of Table 1 reports the estimated adjustment

TABLE 1

JOHANSEN MAXIMUM LIKELIHOOD PROCEDURE

(a) Estimated eigenvalues

USA	UK	Japan	Germany	France	Canada	Italy
0.6532	0.7014	0.6800	0.6845	0.7001	0.8220	0.7257
0.4803	0.5201	0.4309	0.4665	0.5055	0.5071	0.5349
0.3552	0.3106	0.3503	0.3657	0.2628	0.3077	0.3304
0.2896	0.1956	0.2500	0.2160	0.2077	0.2684	0.1920
0.0704	0.1504	0.1182	0.1716	0.1603	0.0862	0.0734

(b) Cointegration likelihood ratio tests based on maximum eigenvalue of the stochastic matrix

H_0	H_1	USA	UK	Japan	Germany	France	Canada	Italy	CV
$r=0$	$r=1$	43.42	49.55	46.72	47.30	49.38	70.77	53.04	37.5
		(38.12)	(43.50)	(41.02)	(41.53)	(43.36)	(62.14)	(46.57)	
$r\leqslant 1$	$r=2$	26.83	30.10	23.11	25.76	28.87	29.01	31.9	31.5
		(23.56)	(26.43)	(20.30)	(22.62)	(25.35)	(25.47)	(27.56)	
$r\leqslant 2$	$r=3$	17.99	15.25	17.68	18.67	12.50	15.08	16.45	25.5
		(15.80)	(13.39)	(15.53)	(16.39)	(10.98)	(13.24)	(14.44)	
$r\leqslant 3$	$r=4$	14.02	8.93	11.79	9.98	9.54	12.81	8.74	19.0
		(12.31)	(7.84)	(10.35)	(8.76)	(8.38)	(11.25)	(7.67)	
$r\leqslant 4$	$r=5$	2.99	6.68	5.16	7.72	7.16	3.69	3.13	12.4
		(2.63)	(5.87)	(4.53)	(6.78)	(6.29)	(3.24)	(2.75)	

(c) Cointegration likelihood ratio tests based on trace eigenvalue of the stochastic matrix

H_0	H_1	USA	UK	Japan	Germany	France	Canada	Italy	CV
$r=0$	$r\geqslant 1$	105.32	110.51	104.50	109.42	107.53	131.40	112.7	37.5
		(92.42)	(97.03)	(91.72)	(96.08)	(94.36)	(115.3)	(98.99)	
$r\leqslant 1$	$r\geqslant 2$	61.84	60.96	57.75	62.12	58.08	60.59	59.71	31.5
		(54.30)	(53.52)	(50.71)	(54.55)	(51.00)	(53.2)	(52.42)	
$r\leqslant 2$	$r\geqslant 3$	35.01	30.86	34.63	36.36	29.21	31.59	28.32	25.5
		(30.74)	(27.09)	(30.41)	(31.93)	(25.65)	(27.73)	(24.87)	
$r\leqslant 3$	$r\geqslant 4$	17.01	15.61	16.95	17.70	16.71	16.51	11.87	19.0
		(14.94)	(13.71)	(14.88)	(15.54)	(14.67)	(14.50)	(10.42)	
$r\leqslant 4$	$r=5$	2.99	6.68	5.16	7.72	7.16	3.69	3.13	12.4
		(2.63)	(5.87)	(4.53)	(6.78)	(6.29)	(3.24)	(2.75)	

(d) Estimated adjustment coefficients q_Y and weak exogeneity tests

	USA	UK	Japan	Germany	France	Canada	Italy
Y	0.34	0.43	0.29	0.11	0.11	0.37	0.35
	[0.04]	[0.06]	[0.03]	[0.01]	[0.02]	[0.06]	[0.07]
LR1	{0.42}	{0.32}	{0.37}	{0.65}	{0.49}	{0.54}	{0.33}

(*continued*)

TABLE 1 *Continued*

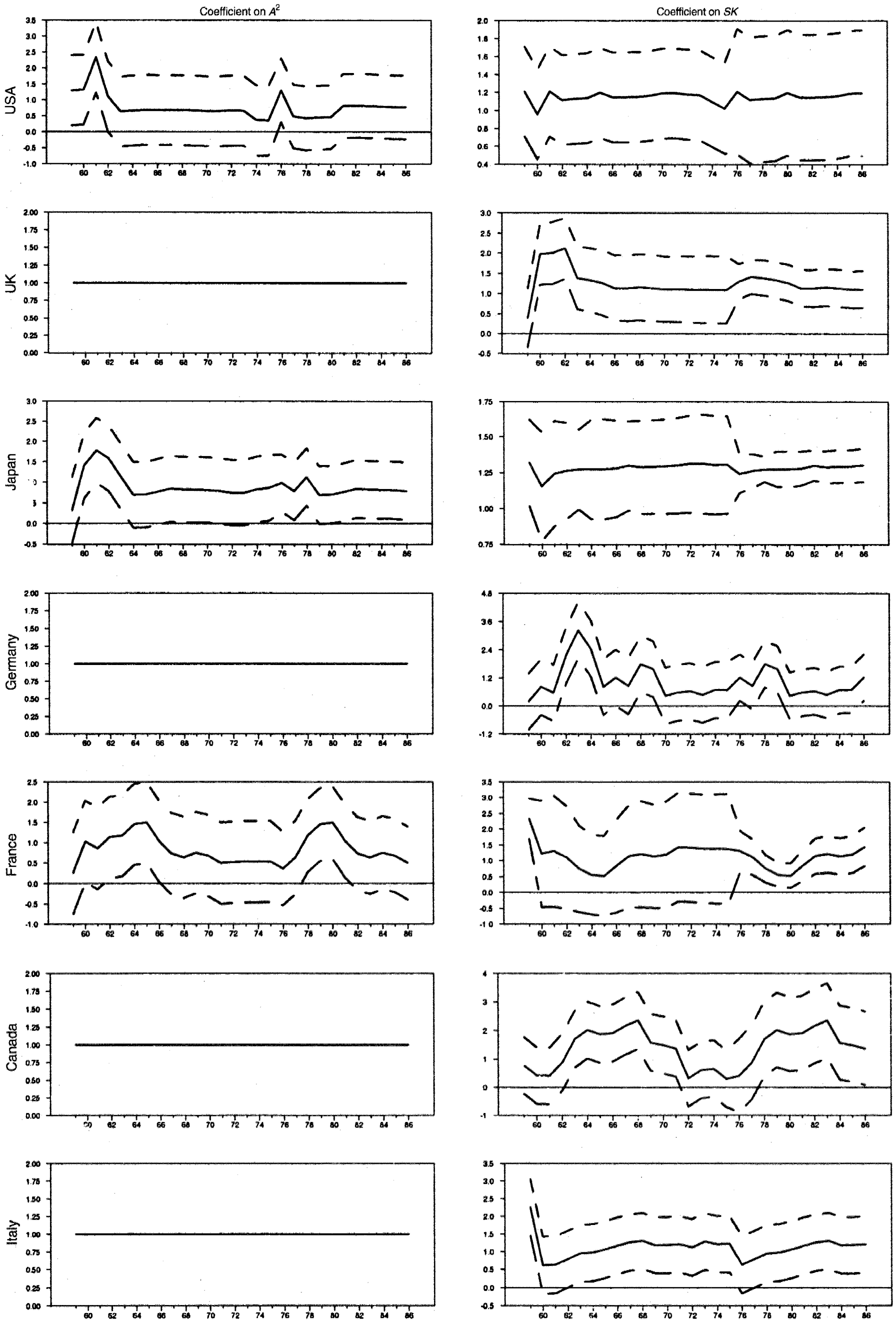
(e) Estimated cointegrating vectors

	USA	UK	Japan	Germany	France	Canada	Italy
A	0.75	1.00	0.79	1.00	0.82	1.00	1.00
	[0.23]	[—]	[0.26]	[—]	[0.17]	[—]	[—]
SK	1.21	1.17	1.32	1.11	1.13	1.18	1.13
	[0.35]	[0.28]	[0.29]	[0.34]	[0.37]	[0.29]	[0.33]
SH	0.76	1.04	0.74	0.90	1.09	0.73	0.74
	[0.13]	[0.26]	[0.21]	[0.24]	[0.16]	[0.15]	[0.20]
Z	−1.93	−2.25	−2.13	−2.09	−2.15	−1.87	−1.92
	[0.41]	[0.47]	[0.57]	[0.49]	[0.62]	[0.38]	[0.56]
LR2	{0.78}	{0.37}	{0.32}	{0.45}	{0.57}	{0.39}	{0.61}

Notes: In panels (b) and (c), H_0 and H_1 denote the null hypothesis and the alternative hypothesis respectively; figures in parentheses are test statistics adjusted for degrees of freedom (Reimers 1992); r denotes the number of cointegrating vectors and CV is the 95% critical value (see Osterwald-Lenum 1992; Johansen 1995, ch. 15, table 15.4). In panels (d) and (e) figures in square brackets denote estimated standard errors, and figures in braces denote p-values. LR1 is the likelihood ratio test statistic for the null hypothesis that the adjustment coefficients associated to A_t, SK_t, SH_t and Z_t are equal to zero when the cointegrating rank equals unity, and is distributed as χ^2 (4) under the null; LR2 denotes the likelihood ratio test statistic for the null hypothesis that the sum of the coefficient on SK_t plus the coefficient on SH_t is equal to the coefficient on Z_t in absolute size, and is distributed as χ^2 (1) under the null.

coefficients for q_Y with their standard errors, in addition to likelihood ratio tests of the joint exclusion restrictions on the four adjustment coefficients associated with A_t, SK_t, SH_t and Z_t, respectively. The results clearly suggest that, while q_Y is very strongly statistically significantly different from zero, weak exogeneity of A_t, SK_t, SH_t and Z_t can be assumed.[9]

In estimating the cointegrating vectors, I also attempted to impose the restriction that the coefficient on A_t equals unity in the cointegrating vector; but this restriction was rejected in three out of seven cases using a likelihood ratio statistic.[10] In panel (e) of Table 1 I then report the estimated cointegrating vectors (after normalizing the coefficient of Y to -1) and their standard errors for each of the G7 countries with the technology homogeneity restriction imposed for the United Kingdom, Germany, Canada and Italy. The estimated cointegrating parameters appear stable over time for each country, as shown visually by the recursive estimates of the unrestricted parameters plotted, together with plus-or-minus twice the corresponding standard errors, in Figure 1. The results are very satisfactory in that the estimated coefficients on SK, SH and Z are correctly signed and the hypothesis that the sum of the coefficient on SK plus the coefficient on SH equals the coefficient on Z (in absolute size) could not be rejected for all countries considered, thereby providing support for the augmented Solow–Swan model. Most interestingly, the estimated coefficients do not differ widely across the G7 countries, yielding sensible implied values of the structural production function parameters. In general, for all countries considered, while human capital appears to play a significant role in the cointegrating regression, the implied estimate of the physical capital accumulation rate is always higher than the implied estimate of the parameter associated with the human capital accumulation rate, i.e. $\alpha > \beta$.



[a]Coefficient on A is restricted for USA, UK, Germany, Canada and Italy.

FIGURE 1. *Continued on p. 411.*

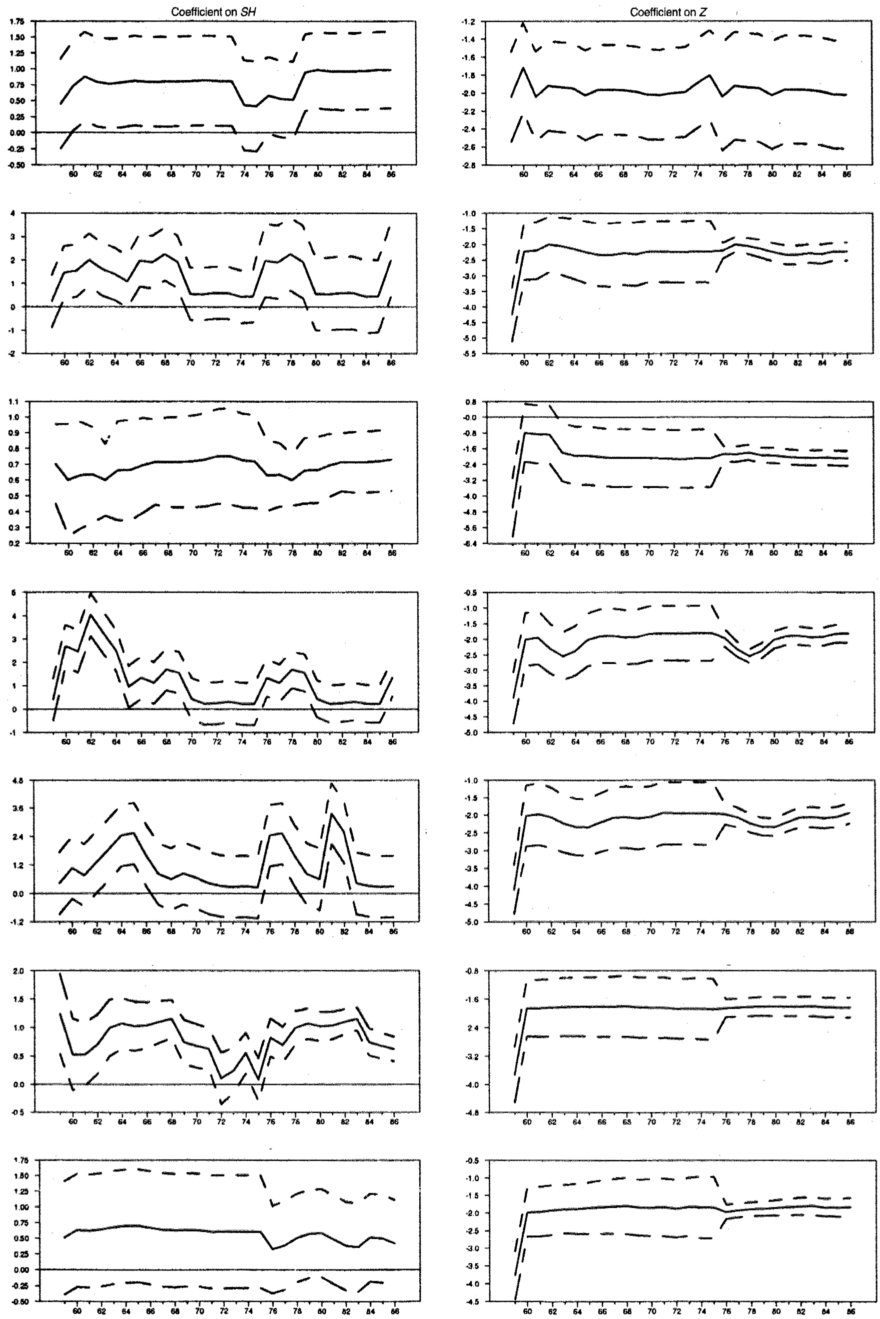


FIGURE 1. Restricted Cointegration Analysis (recursive estimation). Solid line = cointegrating parameter (b); broken lines = $b + 2.0$ times standard error.

The estimated coefficients on SK and SH given in panel (c) of Table 1 imply, for example, that $\alpha = 0.42$ and $\beta = 0.24$ for the United States whereas $\alpha = 0.36$ and $\beta = 0.33$ for the United Kingdom. Thus, notwithstanding the inability to impose the restriction that the coefficient on A_t equals unity in three out of seven cases, the cointegration analysis provides evidence supporting the augmented Solow–Swan model, also suggesting estimated parameters of the production function which are plausible and fairly consistent with my economic priors.

The finding that a cointegrating relationship exists between a set of variables implies that an ECM exists (Granger 1986). Therefore I used the cointegrating residuals—the estimated equilibrium error (u_t)—for each of the seven countries examined, both as the basis for the linearity testing and as the equilibrium correction term in the econometric modelling of the adjustment towards the steady state.

(b) Linear dynamic modelling

Given a joint normally distributed process $\mathbf{X}_t$ with its history denoted $\mathbf{X}^h_{t-1} \equiv (\mathbf{X}_1, \mathbf{X}_2, \ldots, \mathbf{X}_{t-1})$, factorization into conditional and marginal distribution yields:

$$L(\mathbf{X}_t \mid \mathbf{X}^h_{t-1};\ \boldsymbol{\zeta}) = L(Y_t \mid A_t, SK_t, SH_t, Z_t, \mathbf{X}^h_{t-1};\ \boldsymbol{\zeta}_1) \times L(A_t, SK_t, SH_t, Z_t \mid \mathbf{X}^h_{t-1};\ \boldsymbol{\zeta}_2), \tag{9}$$

where $\boldsymbol{\zeta}$, $\boldsymbol{\zeta}_1$ and $\boldsymbol{\zeta}_2$ are vectors of parameters, the first term on the right-hand side of (9) gives the endogenous variable Y_t as a function of $\mathbf{X}^h_{t-1}$, A_t, SK_t, SH_t and Z_t, while the second term on the right-hand side of (9) gives the determination of the exogenous variables A_t, SK_t, SH_t and Z_t as a function of $\mathbf{X}^h_{t-1}$. Then, provided A_t, SK_t, SH_t and Z_t are at least weakly exogenous, the conditional distribution of Y_t may be reparameterized as an ECM where ΔY_t is explained by its lagged values, the equilibrium correction term and simultaneous changes and their lags of the weakly exogenous variables (see Johansen 1995, theorem 8.1).

Nevertheless, given the potential correlation of economic growth across countries, I estimated the linear equilibrium correction models for the G7 countries jointly using feasible generalized least squares (FGLS). By exploiting the contemporaneous correlation of the disturbances across equations, FGLS is expected to yield a significant efficiency gain relative to the estimation of each growth equation individually by ordinary least squares.[11] Then, the linear ECM estimated jointly for each country examined, also allowing for spillover effects in growth dynamics within the G7, is of the form

$$\begin{aligned}\Delta Y_t = k_0 + \rho_{ECM}\left(k_1 + \sum_{i=1}^{7} \rho_i u_{it-1}\right) + \sum_{i=1}^{p} \varphi_{1i}\Delta Y_{t-i} \\ + \sum_{i=0}^{p} \varphi_{2i}\Delta A_{t-i} + \sum_{i=0}^{p} \varphi_{3i}\Delta SK_{t-i} \\ + \sum_{i=0}^{p} \varphi_{4i}\Delta SH_{t-i} + \sum_{i=0}^{p} \varphi_{5i}\Delta Z_{t-i} + \text{innovations},\end{aligned} \tag{10}$$

where k_0 is a constant term, ρ is the speed of adjustment coefficient, and the lag length p was set equal to 2.

The results from the FGLS estimation of parsimonious equilibrium correction models for labour productivity growth for each of the G7 countries, obtained by adopting a conventional general-to-specific procedure as originally suggested by Davidson *et al.* (1978) and Hendry *et al.* (1984), are reported in panel (a) of Table 2. The resulting models appear to be adequate in terms of quite high coefficients of determination and approximately white-noise residuals (see panel (b) of Table 2). For each country $i \in \{1, ..., 7\}$, the speed of adjustment coefficient ρ_i is of particular interest in that it has important implications for the dynamics of the model. In fact, for any given value of the deviations from long-run equilibrium, a large value of the speed of adjustment ρ_i is associated with a large value of the change in labour productivity. If ρ_i is zero, however, the change in labour productivity at time t does not respond at all to the deviation from long-run equilibrium in period $(t-1)$. As shown in Table 2, the estimated coefficient on the equilibrium correction term is found to be statistically significantly different from zero at conventional nominal levels of significance in each case, suggesting—consistent with the results of the cointegration analysis—that there is a long-run relationship of the type implied by the augmented Solow–Swan model. Also, the results provide evidence of spillover effects across the G7, and in particular from the United States to each of the other G7 countries.

We now turn to the empirical examination of potential nonlinearities in the adjustment of labour productivity towards its equilibrium level, as a test of the economic prior that the adjustment towards the steady state in the augmented Solow–Swan model is nonlinear.

(c) Linearity tests and nonlinear equilibrium correction models

I start by providing some evidence of nonlinearity in the estimated equilibrium error u_t from equation (8) using the RESET (Ramsey 1969) test statistic. The RESET test, which is a fairly general misspecification test and uses few degrees of freedom, seems particularly appropriate in the present context, given the small number of observations available. Under the RESET test statistic, the alternative model involves a higher-order polynomial to represent a different functional form; under the null hypothesis, the statistic is distributed as $\chi^2(d)$ with d equal to the number of higher-order terms in the alternative model. In constructing the tests I use the F-statistic form, since it is well known that in finite samples the actual test size of the F approximation may be closer to the nominal significance level than the actual size of the χ^2 approximation (e.g. Harvey 1990, pp. 174–5; Granger and Teräsvirta 1993, pp. 76–9). Table 3 reports the results from executing RESET test statistics obtained constructing a first-order autoregressive model for the cointegrating residuals with an intercept and augmenting that by the squared and cubed fitted values. The results suggest a very strong rejection of the null hypothesis for all countries considered with a p-value of virtually zero.[12]

Given the results from the linearity tests, I estimated nonlinear equilibrium correction models for ΔY, with the estimated equilibrium error retrieved from (8) as the equilibrium correction term. Joint estimation of the seven equilibrium correction models is by multivariate nonlinear least squares (Gallant 1987;

TABLE 2

ESTIMATED PARSIMONIOUS LINEAR EQUILIBRIUM CORRECTION MODELS

(a) Estimated coefficients

	USA	UK	Japan	Germany	France	Canada	Italy
k_0	0.025 (0.010)	0.024 (0.011)	0.027 (0.007)	0.026 (0.005)	0.019 (0.005)	0.024 (0.006)	0.025 (0.008)
ρ_{ECM}	−0.102 (0.035)	−0.155 (0.043)	−0.272 (0.098)	−0.092 (0.024)	−0.115 (0.034)	−0.111 (0.027)	−0.106 (0.024)
k_1	—	0.006 (0.001)	0.073 (0.015)	—	0.008 (0.002)	0.007 (0.002)	0.005 (0.001)
ρ_1	3.190 (1.231)	1.223 (0.457)	0.639 (0.238)	0.643 (0.291)	0.604 (0.175)	1.846 (0.564)	2.389 (1.022)
ρ_2	—	3.203 (1.313)	—	—	—	—	—
ρ_3	—	−0.930 (0.028)	1.235 (0.349)	−0.469 (0.173)	—	0.756 (0.271)	—
ρ_4	—	—	—	1.265 (0.420)	—	—	—
ρ_5	—	—	—	−0.118 (0.036)	1.043 (0.234)	—	—
ρ_6	—	—	—	—	−0.271 (0.058)	3.334 (1.280)	—
ρ_7	—	—	—	—	0.686 (0.228)	—	0.645 (0.210)
φ_{11}	—	—	—	—	−0.446 (0.123)	—	—
φ_{12}	0.289 (0.097)	—	—	0.472 (0.201)	—	—	—
φ_{20}	—	0.289 (0.074)	—	—	—	—	—
φ_{21}	0.347 (0.105)	—	0.387 (0.120)	0.174 (0.062)	0.560 (0.148)	—	0.322 (0.052)
φ_{22}	—	—	—	—	—	0.351 (0.057)	—
φ_{30}	0.405 (0.094)	0.256 (0.048)	0.264 (0.071)	0.456 (0.074)	0.431 (0.156)	0.383 (0.128)	0.402 (0.119)
φ_{31}	0.093 (0.024)	—	—	0.235 (0.066)	—	0.074 (0.022)	0.076 (0.013)
φ_{32}	—	—	—	—	—	—	0.084 (0.021)
φ_{40}	—	—	—	—	—	—	—
φ_{41}	—	0.125 (0.037)	—	—	—	—	—
φ_{42}	—	—	—	—	—	—	—
φ_{50}	—	—	—	—	—	—	—
φ_{51}	—	—	−0.093 (0.024)	−0.260 (0.045)	—	—	—
φ_{52}	—	—	—	—	—	—	—

(b) Diagnostics

	USA	UK	Japan	Germany	France	Canada	Italy
R^2	0.79	0.70	0.68	0.81	0.75	0.74	0.85
JB	{0.43}	{0.57}	{0.32}	{0.60}	{0.78}	{0.41}	{0.84}
DW	1.96	1.82	1.89	2.05	1.90	1.84	1.87
$Q(3)$	{0.39}	{0.47}	{0.75}	{0.38}	{0.85}	{0.71}	{0.29}
STAB	{0.25}	{0.29}	{0.24}	{0.36}	{0.42}	{0.18}	{0.31}
	[t = 1974]	[t = 1973]	[t = 1980]	[t = 1967]	[t = 1970]	[t = 1968]	[t = 1974]

Notes: The estimated coefficients of the linear ECM (10) for each G7 country were obtained by estimating a system of seven equilibrium correction equations by FGLS and applying the general-to-specific procedure, as discussed in the text; figures in parentheses denote estimated standard errors. ρ_i is the coefficient associated with the equilibrium error from country $i \in \{1, \ldots, 7\}$, and countries 1–7 refer to the USA, the UK, Japan, Germany, France, Canada and Italy respectively. R^2 is the proportion of the variation in the dependent variable explained by the model; *JB* is the Jarque–Bera (Jarque and Bera 1980) test for normality and is distributed as $\chi^2(2)$ under the null; $Q(m)$ is the Ljung–Box test for residual autocorrelation up to order m, and is distributed as $\chi^2(m)$ under the null; *DW* is the Durbin–Watson statistic. Figures in braces denote p-values. *STAB* is the smallest p-value from executing a recursive Chow test on the residuals for each equation; figures in square brackets refer to the year corresponding to the smallest p-value.

TABLE 3

RESET TEST STATISTICS ON THE ESTIMATED EQUILIBRIUM ERROR

USA	UK	Japan	Germany	France	Canada	Italy
2.34×10^{-5}	3.50×10^{-6}	2.46×10^{-6}	7.73×10^{-6}	3.82×10^{-5}	9.03×10^{-6}	7.39×10^{-5}

Notes: Figures are p-values from RESET test statistics carried out on the cointegrating residuals retrieved from the static regression (8) for each G7 country and constructed as described in the text.

Gallant and White 1988). Also, in order to select 'good' starting values of the parameters, I used preliminary estimates of linear equilibrium correction models; in the empirical analysis this method proved to be very effective for achieving convergence in the estimation of the nonlinear equilibrium correction models. I then applied a general-to-specific procedure in order to reach parsimonious empirical models. The parsimonious models were obtained by 'testing down' the general system of error correction models with a lag length of 2, imposing exclusion restrictions on the coefficient with the lowest (in absolute size) insignificant t-ratio and re-estimating the system sequentially. I repeated the estimation procedure several times using different alternative sequences and also using different sets of starting values for the parameters, in order to ensure that the results were robust to the specification search rule and that a global minimum was achieved.

After considerable experimentation with various nonlinear possibilities for modelling the adjustment towards long-run equilibrium, the logistic cumulative distribution function suggested by Granger and Swanson (1996, p. 549) was found to work extremely well:

$$h_L(\cdot) = h_L(\hat{\boldsymbol{\lambda}}'\mathbf{U}_{t-1}) = [1 + \exp(-\hat{\boldsymbol{\lambda}}'\mathbf{U}_{t-1})]^{-1} - \frac{1}{2}. \tag{11}$$

Equation (11) also accounts for the possibility that, while cointegration is established in the fundamental static equation of the augmented Solow–Swan model for each of the individual countries, despite long-run separation, the adjustment towards long-run equilibrium is characterized by the equilibrium error from one equation entering other equilibrium correction equations of the same system; that is, I included in $\mathbf{U}_{t-1}$ all the estimated equilibrium errors from each equation.[13] Assuming $h_L(\cdot)$, the nonlinear ECM takes the form

$$\begin{aligned} \Delta Y_t = k_0' + \lambda_{ECM}\left\{[1 + \exp(-\hat{\boldsymbol{\lambda}}'\mathbf{U}_{t-1})]^{-1} - \frac{1}{2}\right\} \\ + \sum_{i=1}^{p} \varphi_{1i}'\Delta Y_{t-i} + \sum_{i=0}^{p} \varphi_{2i}'\Delta A_{t-i} + \sum_{i=0}^{p} \varphi_{3i}'\Delta SK_{t-i} \\ + \sum_{i=0}^{p} \varphi_{4i}'\Delta SH_{t-i} + \sum_{i=0}^{p} \varphi_{5i}'\Delta Z_{t-i} + \text{innovations}. \end{aligned} \tag{12}$$

While the functional form (11) has proved to be very successful in modelling macroeconomic time series in the recent literature on nonlinear equilibrium correction processes (e.g. Granger and Swanson 1996), it is clear that the nonlinear adjustment implied by the generalization of the augmented Solow–Swan model presented above does not provide a compelling rationale for the logistic cumulative distribution function. In fact, my choice to characterize the nonlinear adjustment process in this context using a logistic function is motivated mainly on statistical grounds, as strong empirical evidence in favour of nonlinearity and asymmetry of the type implied by equation (11) has been provided using output, industrial production and other related time series by a number of researchers (notably Granger and Teräsvirta 1993; Teräsvirta 1995; Skalin and Teräsvirta 1999). Inadequacy of the logistic function (11) to describe the adjustment of labour productivity may be expected to show up in the inability to fit a nonlinear ECM with plausible estimates of the parameters, or in the failure to find statistically significant parameters on the logistic function or the equilibrium correction terms.

With this caveat in mind, however, the empirical results discussed below suggest that, in modelling deviations from the steady state, the logistic cumulative distribution function provides a very significant improvement in terms of goodness of fit relative to the best-fitting linear ECM as well as relative to alternative nonlinear equilibrium correction models which also outperformed the best-fitting linear ECM.

Also, one plausible interpretation of the logistic function may be the following. Suppose that the linear equilibrium correction term is a tangent of the logistic function passing through the point of inflection of the logistic function itself. The estimated model employed here then indicates that the strength of attraction approaches a constant as one moves out of the equilibrium, while the equilibrium correction would indicate that the strength of attraction increases linearly in the same situation. In addition, the finding that a logistic function well characterizes the adjustment towards the stochastic steady state may also be interpreted as evidence of business cycle asymmetries. In that sense, the logistic function describes a situation where the contraction and expansion phases of an economy display different dynamics (see the references in Granger and Teräsvirta 1993, pp. 141–7; Skalin and Teräsvirta 1999).

The results from estimating nonlinear equilibrium correction models with function (11) describing the nonlinear adjustment, reported in panel (a) of Table 4, indicate that the estimated nonlinear equilibrium correction models have similar features across countries, displaying a very strongly significant association between ΔY and ΔSK, and with some of the equilibrium errors from economies of the G7 found significant in other equilibrium correction models examined. As one might expect, the US equilibrium error is found to be significant in each of the estimated nonlinear equilibrium correction models, suggesting that spillover effects in growth dynamics from the US economy are very strongly significant within the G7. In contrast, spillover effects from other G7 countries are not found to have any statistically significant effect in the United States. Also, while the equations for Japan and Italy display statistically significant spillover effects only from the United States, the remaining four G7 countries appear to have more complex dynamics, in that labour productivity changes respond to three (for the United Kingdom and Canada) or four (for Germany and France) estimated G7 equilibrium errors. For each country, however, as one might expect, labour productivity movements respond most

TABLE 4

ESTIMATED PARSIMONIOUS NONLINEAR EQUILIBRIUM CORRECTION MODELS

(a) Estimated coefficients

	USA	UK	Japan	Germany	France	Canada	Italy
k'_0	0.026 (0.009)	0.025 (0.007)	0.030 (0.011)	0.027 (0.009)	0.024 (0.003)	0.020 (0.004)	0.023 (0.008)
λ_{ECM}	0.005 (0.001)	0.011 (0.003)	0.025 (0.005)	0.011 (0.002)	0.008 (0.002)	0.012 (0.003)	0.018 (0.004)
λ_0	0.012 (0.003)	0.006 (0.001)	0.008 (0.002)	0.009 (0.003)	0.011 (0.003)	0.006 (0.002)	0.007 (0.002)
λ_1	−2.671 (0.783)	−3.742 (0.982)	−0.846 (0.127)	−4.219 (1.293)	−2.124 (0.982)	−2.247 (0.859)	−2.360 (0.839)
λ_2	—	−6.243 (1.641)	—	—	—	—	—
λ_3	—	1.610 (0.692)	−3.461 (1.027)	1.245 (0.325)	—	—	—
λ_4	—	—	—	−5.973 (1.012)	—	−1.293 (0.420)	—
λ_5	—	—	—	—	−4.536 (0.762)	—	—
λ_6	—	—	—	1.239 (0.285)	−1.239 (0.430)	−4.352 (1.089)	—
λ_7	—	—	—	—	−2.653 (0.892)	—	−3.756 (1.105)
φ'_{11}	0.381 (0.078)	—	—	—	0.423 (0.112)	—	—
φ'_{12}	0.288 (0.083)	—	0.258 (0.102)	0.649 (0.221)	—	—	—
φ'_{20}	—	—	—	0.732 (0.127)	0.328 (0.126)	—	0.562 (0.121)
φ'_{21}	0.543 (0.153)	0.375 (0.081)	—	—	—	0.251 (0.072)	
φ'_{22}	—	—	0.172 (0.036)	—	—	—	—
φ'_{30}	0.329 (0.062)	0.286 (0.095)	0.283 (0.052)	0.467 (0.116)	0.364 (0.073)	0.362 (0.103)	0.402 (0.127)
φ'_{31}	0.073 (0.012)	—	—	0.217 (0.103)	—	—	0.113 (0.029)
φ'_{32}	—	—	—	—	—	—	—
φ'_{40}	—	—	—	—	0.302 (0.082)	—	0.087 (0.022)
φ'_{41}	—	0.102 (0.034)	—	—	—	0.118 (0.051)	—
φ'_{42}	—	—	—	—	—	0.102 (0.024)	—
φ'_{50}	—	—	−0.192 (0.045)	—	—	—	—
φ'_{51}	—	—	—	−0.221 (0.093)	—	—	—
φ'_{52}	—	—	—	—	—	—	—

(b) Diagnostics

	USA	UK	Japan	Germany	France	Canada	Italy
R^2	0.86	0.82	0.84	0.92	0.90	0.88	0.94
JB	{0.62}	{0.35}	{0.45}	{0.63}	{0.54}	{0.75}	{0.58}
DW	1.90	1.79	1.85	1.83	2.07	1.89	2.08
Q(3)	{0.76}	{0.70}	{0.62}	{0.51}	{0.59}	{0.77}	{0.60}
STAB	{0.34}	{0.25}	{0.36}	{0.22}	{0.37}	{0.21}	{0.19}
	[t = 1973]	[t = 1977]	[t = 1980]	[t = 1967]	[t = 1974]	[t = 1972]	[t = 1974]

Notes: The estimated coefficients of the nonlinear ECM(12) for each G7 country were obtained by estimating a system of seven nonlinear equilibrium correction equations by multivariate nonlinear least squares and applying the general-to-specific procedure, as discussed in the text; figures in parentheses denote estimated standard errors. The vector $U_{t-1} = [1 - u_{1,t-1}\ u_{2,t-1} \ldots u_{7,t-1}]'$, where $u_{i,t-1}$ ($i = 1, 2, \ldots, 7$) is the estimated error term from the cointegrating regression (8) for each of the G7 countries in the following order: USA, UK, Germany, France, Japan, Canada, Italy; R^2 is the proportion of the variation in the dependent variable explained by the model; *JB* is the Jarque–Bera (Jarque and Bera 1980) test for normality and is distributed as $\chi^2(2)$ under the null; $Q(m)$ is the Ljung–Box test for residual autocorrelation up to order m; DW is the Durbin–Watson statistic. Figures in braces denote p-values. *STAB* is the smallest p-value from executing a recursive Chow-type test on the residuals for each equation; figures in square brackets refer to the year corresponding to the smallest p-value.

strongly to the deviations from the steady state in the domestic country. Moreover, for each model the estimated parameter on the logistic cumulative distribution function is found to be very strongly significantly different from zero, clearly suggesting that labour productivity responds at time t to deviations from long-run equilibrium at time $(t-1)$ in a nonlinear fashion, and therefore exhibits nonlinear convergence towards the steady state.

A battery of diagnostics is reported in panel (b) of Table 4, including a Jarque–Bera test for normality, a Ljung–Box test for residual serial correlation and a Durbin–Watson statistic, none of which was found significant at conventional nominal levels of significance.[14] I also tested for empirical stability of the model by constructing a Chow-type recursive test for each nonlinear ECM. The results suggest no structural break in the residuals, with the smallest p-value (reported in the last column of panel (b)) reasonably larger than the conventional 5%.

Goodness-of-fit statistics are also found to be very satisfactory: the coefficients of determination, always significantly higher than the coefficients of determination of the corresponding linear equilibrium correction models reported in Table 2, are in the range 0.82–0.94. Nevertheless, I compared the performance of the nonlinear ECM (12) not only relative to the linear ECM (10) but also relative to two alternative nonlinear equilibrium correction models which were found to outperform the linear ECM during the experimentation, leading to the selection of the nonlinear ECM (12). The first of these nonlinear models is of the form (12) but the nonlinear function employed to model the adjustment towards equilibrium is the exponential function

$$(13) \qquad h_E(\cdot) = [1 - \exp(-\hat{\boldsymbol{\lambda}}'\mathbf{U}_{t-1}^2)].$$

The exponential function (13) is bounded between zero and unity, with $h_E(x) = 0$ and $\lim_{x \to \pm\infty} h_E(x) = 1$, and is symmetric around zero. These properties would generate, in the present context, symmetric adjustment of labour productivity movements for deviations above and below the steady state, implying faster adjustment the larger the deviation from the steady state.

The second nonlinear ECM considered is an ECM where the equilibrium correction term u_t is replaced, for each country, by u_t^+ and u_t^-, with

$$(14) \qquad \begin{cases} u_t^+ = u_t & \text{if} \quad u_t \geqslant 0,\ u_t = 0 \text{ otherwise} \\ u_t^- = u_t - u_t^+. \end{cases}$$

Hence this nonlinear ECM also allows a varying strength of attraction to the steady state, but the attractor is assumed to be stronger on one side than on the other, so that u_t^+ may have a different coefficient from u_t^- (see Granger and Lee 1989).

In order to compare the goodness of fit of the nonlinear ECM (12) to the linear ECM (10) and the other nonlinear equilibrium corrections described above, I calculated the ratio of the R^2, the residual variance (RV), the AIC and the SIC from each of the estimated nonlinear ECM (12) reported in Table 4 to the corresponding measure for each of the three alternative models considered. The results, reported in Table 5, show that, for each country examined, the estimated nonlinear ECM (12) largely outperforms the best alternative linear

TABLE 5
RELATIVE GOODNESS OF FIT

	USA	UK	Japan	Germany	France	Canada	Italy
Nlin1 versus Linear							
R^2 ratio	1.09	1.17	1.23	1.13	1.20	1.19	1.10
RV ratio	0.82	0.80	0.73	0.81	0.77	0.75	0.79
AIC ratio	1.13	1.22	1.24	1.19	1.26	1.20	1.18
SIC ratio	1.14	1.24	1.23	1.20	1.24	1.21	1.18
Nlin1 versus Nlin2							
R^2 ratio	1.06	1.12	1.19	1.10	1.09	1.11	1.07
RV ratio	0.85	0.83	0.78	0.84	0.80	0.81	0.82
AIC ratio	1.10	1.15	1.19	1.16	1.21	1.19	1.15
SIC ratio	1.12	1.17	1.17	1.16	1.20	1.18	1.13
Nlin1 versus Nlin3							
R^2 ratio	1.07	1.14	1.20	1.11	1.11	1.14	1.10
RV ratio	0.84	0.82	0.76	0.83	0.79	0.77	0.80
AIC ratio	1.11	1.18	1.22	1.17	1.23	1.19	1.17
SIC ratio	1.12	1.18	1.23	1.18	1.22	1.20	1.16

Notes: *Linear, Nlin1, Nlin2* and *Nlin3* refer to the estimated linear equilibrium correction models reported in Table 2, the estimated nonlinear equilibrium correction models with the adjustment characterized by a logistic cumulative distribution function reported in Table 4, the estimated nonlinear equilibrium correction models with the adjustment characterized by the exponential function (13) and the estimated nonlinear equilibrium correction models which allow for different strength of attraction according to the specification (14), respectively. The R^2 ratio, the *RV* (residual variance) ratio, the *AIC* and the *SIC* ratios are the ratios of the R^2, the residual variance, the *AIC* and the *SIC*, respectively, from each country's estimated nonlinear ECM with logistic cumulative distribution function reported in Table 4 (i.e. *Nlin1*) to the corresponding goodness-of-fit measure obtained for the alternative model.

model, leading to a substantial reduction—up to 27% for Japan—of the residual variance. Moreover, both the nonlinear ECM with an exponential function describing the adjustment towards equilibrium and the nonlinear ECM allowing for different strength of attraction appear to outperform the linear ECM, although they are outperformed in goodness of fit by the nonlinear ECM (12) on the basis of the four goodness-of-fit measures employed.

(d) Summing up the empirical results

Overall, the results reported in this section suggest that significant nonlinearity is present in the error representing the deviation from long-run equilibrium in the static equation implied by the augmented Solow–Swan model, consistent with my priors. Also, the nonlinear adjustment towards the steady state could be modelled very satisfactorily using a number of different alternative nonlinear specifications. The preferred specification was, however, a logistic cumulative distribution function, which yielded plausible estimates of the parameters and insignificant diagnostics, enabling us to identify a potential data-generating mechanism of economic growth for each of the countries considered and significantly outperforming the alternative linear and nonlinear models. Finally, the equilibrium errors from countries of the G7 often entered the equilibrium correction equations of other countries of the group, suggesting the presence of significant spillover effects in growth dynamics.

V. Concluding Remarks

Traditional growth theory in line with the augmented Solow–Swan model implies a long-run relationship between labour productivity and the determinants of the steady state, while the steady state displays cross-country heterogeneity and varies over time, and the adjustment process towards equilibrium is consistent with a nonlinear ECM. Using postwar data for the G7 countries, I was able to establish cointegration in the static regression predicted by the augmented Solow–Swan model, with correctly signed and plausible implied estimates of the production function parameters. My results clearly suggest that technology, capital accumulation rate and labour force growth generate a stochastic long-run equilibrium level of labour productivity. Most tellingly, I provide strong empirical evidence that the adjustment towards the steady state occurs in a nonlinear fashion and that the growth path of each of the G7 countries may be satisfactorily modelled as a nonlinear ECM with the adjustment towards long-run equilibrium governed by a logistic cumulative distribution function.

While these results raise interesting issues in the relevant literature, they should be seen as tentative evidence, and further empirical work might be carried out towards establishing the robustness of my conclusions and applying the methods employed in this paper to data for developing countries. Also, empirical growth models may be enriched by considering other factors, such as proxies for education policies, corruption and political instability (e.g. Alesina and Perotti 1996; Blomberg 1996), variables related to the medium-term macroeconomic performance of a country that may affect cross-country differences in economic growth patterns (see e.g. Andres *et al.* 1996), or proxies for stabilization policies that may have persistent effects on economic growth (see Blackburn 1999).

Finally, although the empirical results reported in this paper are fairly satisfactory, a deeper understanding of the true unknown data-generating mechanism of growth dynamics may be gained by experimenting with alternative stochastic nonlinear models and, perhaps more importantly, by strengthening the link between economic theory and its empirical counterpart. These issues remain immediate avenues for future research.

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NOTES

1. Notably see Lee *et al.* (1997, 1998); see also Quah (1993, 1996a), Islam (1995), Cellini (1997), Lee (1998), Durlauf and Quah (1999); see also Crafts and Toniolo (1996) and the papers therein.
2. Also, the stochastic Solow–Swan model leads to a notion of convergence which is different from the definitions of convergence traditionally used in the relevant literature (Lee *et al.* 1997, 1998); see also the related studies by Quah (1993, 1996a), Bernard and Durlauf (1995), Islam (1995), Lee (1998) and Durlauf and Quah (1999).
3. For example, Lee *et al.* (1997) provide evidence that steady-state growth rates are likely to differ across countries and that the growth of technology has been higher in OECD countries, with smaller (insignificantly different from zero) dispersion, compared with the world as a whole.
4. Empirical evidence in favour of nonlinearities in output and industrial production has been provided by a number of researchers; see e.g. Teräsvirta (1995) and Skalin and Teräsvirta (1999).
5. For an excellent treatment of the concepts and techniques discussed in this section, see Granger and Teräsvirta (1993), Granger and Swanson (1996) and Granger and Haldrup (1997); see also Escribano (1986), Granger and Lee (1989) and Escribano and Mira (1996). These references also provide extensive discussion on the relevant specification, estimation and evaluation procedures which are, therefore, not repeated here.
6. Although data on some of the variables required for the present application are available from other sources for the countries examined, I prefer to use the Summer–Heston data-set since it allows direct comparability with much relevant literature on growth empirics, while being of very high quality for the G7 countries. In this way, the problems typically encountered in estimating growth regressions with potential measurement errors and dubious data—recently rigorously addressed by, *inter alios*, Leamer and Taylor (1999)—are minimized.
7. Zvi Griliches and Ron Smith are gratefully acknowledged for their suggestions and guidance in the construction of the R&D stock measure. The main reason for using the R&D stock is that, although R&D does not capture everything covered by the notion of technology, R&D-based indicators are the most widely used proxies in this context and, most tellingly, R&D activities have grown in importance as sources of technology (Patel and Pavitt 1995; Griliches 1992, 1994, 1995, 1998; Geroski 1995; Jones 1995; Temple 1999).
8. Given the convincing recent empirical evidence provided by Lee *et al.* (1997) suggesting that differences in the technological progress rate g are not statistically significant within the OECD, I assumed that g equals 0.027—the estimated value of g provided by Lee *et al.* (1997) for 22 OECD countries—for each G7 country. Nevertheless, allowing g to vary across countries and setting it equal to the estimated mean value of ΔA_t yielded estimates for g of between 0.024 and 0.029, and the restriction that g is the same across the G7 countries could not be rejected with a p-value of 0.45 using a likelihood ratio test. In addition, I tested the hypothesis that there is no structural break in the mean of ΔA_t using both a recursive Chow test statistic and a Vogelsang (1997) test statistic which allows for the break-point to be chosen endogenously, but I could not reject the null hypothesis of no structural break, suggesting that the assumption of a constant g in the empirical analysis may be tenable in the present application. I adopted the conventional assumption that the depreciation rate δ, for which no good data are available, equals 0.03 for all countries examined (see Mankiw *et al.* 1992; Barro and Sala-i-Martin 1995; Aghion and Howitt 1998; Jones 1998).
9. Execution of weak exogeneity tests on each of A_t, SK_t, SH_t and Z_t separately (i.e. testing exclusion restrictions on the individual adjustment coefficients) also indicated non-rejection of the null hypothesis of weak exogeneity in every case. In addition, I also tested for strong exogeneity by assuming weak exogeneity and applying Granger non-causality tests for feedbacks from ΔY_t. Since Y and the four explanatory variables are cointegrated, Granger non-causality can also be tested within the complete VAR model in $I(1)$ space, and the tests for excluding coefficients of lagged Y_t have a limiting χ^2 distribution as shown, for example, in Watson (1994). The results (not reported to conserve space, but available from the author on request) indicated rejection of strong exogeneity for each of the variables in the VAR.
10. The restriction of a unity coefficient on A_t could not be rejected for the UK, Germany, Canada and Italy with p-values equal to 0.32, 0.25, 0.46 and 0.39 respectively, while it was rejected for the USA, Japan and France with p-values equal to 0.04, 0.03 and 0.04 respectively—these likelihood ratio test statistics are obviously distributed as $\chi^2(1)$. I also tested exclusion

restrictions on the deterministic trend in the cointegration space, but for each country I was able to reject the hypothesis that the trend is not statistically significant at conventional nominal significance levels.

11. A likelihood ratio test statistic of the null hypothesis that the variance–covariance matrix obtaining from the changes in labour productivity in each of the G7 countries is diagonal suggested very strong rejection of the null hypothesis (with a p-value of virtually zero), implying that FGLS estimation may yield a large efficiency gain.
12. Higher-order autoregressive terms for the cointegrating residuals were found to be statistically insignificant at conventional nominal levels of significance. Moreover, using a second- or a third-order autoregressive model yielded qualitatively identical results. Also, when a first-order model is used, the RESET test is equivalent to the linearity test derived by Luukkonen *et al.* (1988).
13. Hence, $\mathbf{U}_{t-1}=[1u_{1,t-1}\ u_{2,t-1}\ \dots\ u_{7,t-1}]'$, where $u_{i,t-1}$ $(i=1,2,\dots,7)$ is the estimated equilibrium correction term from (8) for each of the G7 countries in the order: USA, UK, Japan, Germany, France, Canada, Italy. Clearly, $\boldsymbol{\lambda}'$ is a vector of parameters $-\lambda'=[\lambda_0\lambda_1\lambda_2\ \dots\ \lambda_7]$.
14. Note, however, that the Ljung–Box statistic does not have its standard asymptotic χ^2 distribution since the residuals come from a nonlinear model (e.g. Godfrey 1988, pp. 117–18; Eitrheim and Teräsvirta 1996)—this is probably the case also for the Durbin–Watson statistic. Nevertheless, given the very large p-values obtained from assuming a χ^2 distribution for the Ljung–Box test, unless the unknown distribution of the test deviates extremely from the χ^2 distribution in the present case, it may seem unlikely that the null hypothesis of no serial correlation could be rejected at conventional nominal levels of significance.

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