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Empirical Evidence for Alternative Growth Models: Time Series Results

By

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I. Growth Theory: Old and New

Explaining the large differences in per capita incomes across the world is one of the most demanding tasks for economists. This task requires a theory of economic development to identify the key variables. For years, the neoclassical model of growth with constant returns to scale [Solow, 1956] has served as a workhorse for empirical analysis. This model has been extended to include more than two factors of production, and to account for changes in the quality of the factors of production. From a theoretical point of view, the most important contributions have been the inclusion of vintage effects of the stock of fixed capital[1] [Solow, 1960], and the inclusion of human capital effects [Schultz, 1961; Becker, 1964]. Jorgenson and Griliches [1967] demonstrated that the unexplained residual derived from the simple two-factor model is reduced to a trivial magnitude if changes in the quality of capital are corrected in the same way as changes in the quality of labor. Taken together, the extensions of the simple model have provided the theoretical framework for the growth accounting literature pioneered by Denison [1962].[2] This research program has flourished and provided many important insights on the sources of economic growth.[3]

Remark: I thank Joachim Fels, Ulrich Hiemenz, and an anonymous referee for helpful comments.

[1] See Denison [1964] for an assessment of the empirical relevance of this concept for the United States.

[2] See Blaug [1970] for a critique of the growth accounting framework, especially with respect to human capital.

[3] See Maddison [1987] for a survey of the literature.

Still, this neoclassical model of economic development suffers from two shortcomings: its inability to account for observed persistent differences of per capita incomes across countries and its strong and evidently counterfactual prediction that international trade should induce catching up toward convergence in capital intensities and factor prices [Lucas, 1988, p. 17]. Furthermore, actual trade patterns reveal that nearly half the world's trade occurs between industrial countries with similar relative factor endowments. Hence, the stylized empirical facts that are difficult to explain within a constant-returns-to-scale model are the apparent lack of international capital flows to poor countries, and the volume (and the composition) of trade between countries with strikingly similar factor endowments.

Therefore, some economists claim that there is a need for alternative theories of trade and growth.[4] This is not to say that before the "new" models arrived, economists had not been able to explain why single developing countries, say in Africa, had not caught up in terms of per capita incomes, and why others, say in Southeast Asia, had. Yet, the problem is that the key variables that are held to be responsible for the observed pattern of development are not identified by the traditional constant-returns-to-scale model.

The new approaches may be labelled learning-by-doing- and R&D-models as a reminiscence of the seminal contributions by Arrow [1962] and Uzawa [1965].[5] The common idea underlying the new approaches and their forerunners is that knowledge is taken as an input in the production function. This renders increasing returns to scale quasi automatically, even if each firm is assumed to produce at constant returns to scale. The reason is that a doubling of all (tangible) inputs will double output in an environment with a *constant* level of knowledge; here, the economy-wide level of knowledge increases because knowledge is taken to be a factor of production. If the increasing returns are external to the firm, a competitive equilibrium exists, since only capital and labor are paid their marginal products, while knowledge is treated as a public good.

The crucial difference between the early increasing-returns-to-scale models and especially Romer's analysis [1986] is the assumption that knowledge displays *increasing* marginal productivity. That is, although new knowledge is assumed to be produced under diminishing returns (as in the early models), the production of goods with new

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[4] See, e.g., Romer [1986; 1990], Lucas [1988], and Grossman and Helpman [1991].
[5] See Sala-i-Martin [1990] and Shaw [1992] for surveys of the literature.

knowledge is assumed to result in increasing returns. This difference in assumptions has a decisive consequence for the steady-state solution: in the early models, the knowledge externality is reflected as a level effect only, in the new approaches, it renders a growth effect. Romer [1986] demonstrates that the key variables externalities, increasing returns in the production of output, and decreasing returns in the production of knowledge are consistent with a competitive equilibrium. The implication is that the new models can explain why per capita incomes may grow without bounds and why, contrary to conventional wisdom, the rate of return to capital may actually increase with rising incomes. Hence, the new models offer an alternative to explain observed income differentials and trade patterns across the world by identifying the key variables within the model.

Thus, the question arises whether catching-up and convergence as suggested by the constant-returns-to-scale model, or persistent differences in per capita incomes as a possible, though not necessary outcome of the new models, are adequate descriptions of the real world. One of the first attempts to empirically discriminate between the old and the new approach was made by Romer [1989] by using cross-country data. His findings in favor of the new approach were successfully rejected by Mankiw et al. [1992], using the same set of data but a constant-returns-to-scale model with explicitly accounting for human capital. Their results and the results in Barro [1991] support the hypothesis of conditional convergence, i.e., poor countries tend to grow faster than rich countries holding constant the determinants of the steady state. Still, Quah [forthcoming] in turn questions this interpretation. He shows that allowing for stochastically time-varying permanent or growth components, economies across the world seem to be converging to a distribution where many remain wealthy and many remain poor. So the profession is left with a full circle of cross-section results.

Recent time series analyses seem to provide more clear-cut answers. They seem to support the new models, especially because the evidence refers to a small number of industrialized countries which are obviously not too different with respect to their discount rates, their population growth, their production technologies, and their institutional framework. Therefore, here at least the concept of conditional convergence should apply. De Long [1991], for instance, finds a strong association between machinery investment shares in GDP and GDP per capita growth over the past century for five industrialized countries. At first sight, this result appears to be inconsistent with the steady-state solution of the traditional model, but not with the possible outcomes of the new models. Bernard and Durlauf [1991] find substantial persistence in the estimated time-series representation of cross-country output deviations which implies no catching-up and no convergence of per capita incomes. This finding, too, can be interpreted as corroborating the new models.

In this paper, I show that the time series evidence does not uniformly support the new models. Using alternative econometric models, I demonstrate that it is not possible to empirically discriminate between the new and the traditional growth models with the data at hand. Theoretical considerations suggest that the results which favor the acceptance of the new models may systematically suffer from a small sample bias. As a consequence, less restrictive alternative econometric specifications lead to results that are more favorable for the traditional model. Hence, a full circle of results is achieved with time series data as well.

II. Alternative Econometric Approaches to Testing Growth Theories

Consider a Cobb-Douglas production function with three input factors of the form [Mankiw et al., 1992]

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

where Y is output, K physical capital, H the stock of human capital, L labor, and A the level of technology, with $\alpha + \beta < 1$, which implies decreasing returns to each input alone and guarantees the existence of a steady state. L and A are assumed to grow exogenously at rates n g, and the number of effective units of labor, $A_t L_t$, grows at rate $n + g$. A constant fraction of output, s, is invested, and the rate of depreciation of both the physical and the human capital is δ. For $\beta = 0$, the above model reduces to the traditional two-factor growth model. It becomes a "new" growth model for $\alpha + \beta = 1$, which implies that there is no steady state to which the model economy converges, since exogenous shocks have persistent effects within the latter model.

The off-steady state properties of the constant-returns-to-scale model can be derived by approximating the steady-state level of output per effective worker, y^*. This leads to a formula for the speed of convergence to the steady state, λ, which is given by [Mankiw et al., ibid.]

$$\frac{d \ln y_t}{dt} = \lambda (\ln y^* - \ln y_t), \tag{2}$$

where $\lambda = (n + g + \delta)(1 - \alpha - \beta)$.

Now it is easy to see that the two-factor model predicts a faster speed of convergence than the extended three-factor model. For $\alpha = \beta = 1/3$, for instance, the model without human capital ($\beta = 0$) predicts a speed of convergence that is two times faster than in the extended model. Assuming $(n + g + \delta) = 0.06$, the halfway time to steady state is about 35 years for the extended model,[6] and about 17 years for the two-factor model.[7] Hence, for testing the steady-state prediction of the traditional model, one has to consider very long time periods. With the two worldwide oil price shocks, the time span since the second world war may mainly reflect off-steady state behavior, and even the whole time span since the turn of the century may not provide sufficient steady-state information, given the additional shocks of the first world war and the Great Depression.

With an increasing-returns-to-scale model as the underlying theoretical framework, one would ignore the distinction between steady-state and non-steady-state behavior and instead would ask whether permanent movements in the per capita income of a certain country are associated with permanent movements in the per capita incomes of other countries. An empirical rejection of this hypothesis is evidence against the constant-returns-to-scale model, since such a result would imply that the per capita incomes of different countries seem to follow independent random walks (possibly with a deterministic trend component) and, therefore, do not converge.

The recently introduced concept of cointegration analysis [Engle and Granger, 1987] provides a relatively simple time-series framework for testing the hypothesis that there are stable long-run relationships between the per capita incomes of relatively poor and rich countries. The existence of such a relationship is a necessary, though not sufficient condition for a catching-up process. However, cointegration tests will provide unbiased estimates for large samples only. Put differently, since cointegration tests are designed to estimate stable *long-run* equilibria, the data at hand have to cover a time span long enough to provide sufficient long-run information. The dilemma for empirical research is that a given set of data may either be interpreted as reflecting cointegrating relationships or off-steady-state behavior. These

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6 This theoretically predicted speed of convergence is confirmed by cross-section analyses for international output movements [Mankiw et al., 1992], regional output movements within European economies [Barro and Sala-i-Martin, 1991], and regional output movements within the United States [Barro and Sala-i-Martin, 1992].

7 Referring to estimates from Denison [1962], Lucas [1988] estimates a halfway time of about 11 years for the two-factor model.

alternative interpretations, however, lead to alternative econometric model specifications, null hypotheses, and testing procedures.

Testing for stable long-run relationships between the per capita incomes of different countries requires a relatively flexible econometric specification. First of all, the functional form of the empirical model has to be considered. For instance, think of Y_t^p as representing the log of per capita income in a relatively poor country, and of Y_t^{us} as representing the log per capita income of a rich country (United States) to which the initially poor country is assumed to catch up and eventually to converge. Then, a linear regression of Y_t^p on Y_t^{us} and a constant is not an appropriate framework, since in this case the estimated parameter value of Y_t^{us} is a constant elasticity.[8] This specification exludes convergence by definition, because it does not allow for a gradual adjustment process which may lead to common (conditional) steady state levels of per capita incomes.

A less restrictive specification which could be used for the convergence regression was first suggested by Working [1943] and popularized in applied demand analysis by Deaton and Muellbauer [1980]. This specification reads

$$y_t^i = c + \theta\, Y_t^{US} + z_t, \tag{3}$$

where y_t^i is the per capita GDP of the initially poor country divided by the per capita GDP of the initially rich country, Y_t^{us} is the log per capita income of the initially rich country, c and θ are parameters, and z_t is an error term.[9] The parameter θ is used to compute the "expenditure" elasticity η_i, the elasticity of per capita GDP in the relatively poor country with respect to the per capita GDP in the rich country:

$$\eta_{\mathrm{i}} = 1 + \theta/\bar{y}_t^i, \tag{4}$$

where $\bar{y}_t^i$ equals $1/\mathrm{T}\ \Sigma\, y_t^i$.

Equation (3) has a straightforward interpretation with respect to catching-up and convergence. A statistically significant positive coefficient indicates that the relatively poor country is catching up. It follows from (4) that the implication of such a finding is a variable

[8] See Bernard and Durlauf [1991], who use this double-log specification to test for catching-up and convergence.

[9] In terms of demand analysis, y^i is the expenditure share of good i, and Y is the log of total consumption expenditures.

elasticity which asymptotically approaches 1 as the catching-up proceeds. If the regression constant c in (1) is found to be not statistically different from zero, then a variable elasticity approaching 1 means that the hypothesis of convergence in terms of a common per capita income can not be rejected. Alternatively, a statistically significant positive constant means a steady state level of per capita income in the poor country which is *lower* than in the rich country, and a statistically significant negative constant means a steady state level of per capita income which is *higher* than in the rich country (conditional convergence).

Estimation of (1) by OLS will deliver unbiased estimates of the parameters c and θ as long as this equation represents a cointegrating relationship and no small sample bias is present. Testing whether (3) actually describes a cointegrating relationship by one of the procedures suggested by Phillips and Ouliaris [1990] or by the alternative procedure suggested by Schmidt and Phillips [1992] involves an analysis of the residual z_t. The hypothesis of cointegration is rejected if z_t contains a unit root, which is observationally equivalent to a high degree of autocorrelation [Cochrane, 1991]. Still, autocorrelated errors also may indicate a misspecified functional form or a dynamic misspecification. Therefore, a misspecified functional form as well as a dynamic misspecification may lead to an unjustified rejection of a cointegrating relationship. The alternative to the cointegration approach is to begin the analysis with a general dynamic model, to employ some diagnostic checks, and then to proceed with parameter estimation.

Consider the autoregressive-distributed lag model (AD 1,1) of the form

$$y_t^i = \beta_0 + \beta_1 Y_t^{us} + \beta_2 Y_{t-1}^{us} + \beta_3 y_{t-1}^i + e_t, \tag{5}$$

where e_t is an independent error term with mean zero and common variance.

This model is fairly general in that it encompasses nine alternative dynamic models as special cases [Hendry et al., 1984]. If it is not rejected by a misspecification test, one can be reasonably confident that the long-run parameters have good statistical properties. For the present analysis it is unneccessary to achieve parsimony in the short-run dynamics by subsequent re-estimation, since the focus here is on the long-run parameters.

Wickens and Breusch [1988] suggest that (5) should be transformed in such a way as to allow point estimates of the long-run parameters and their standard errors. This specification reads[10]

$$y_t^i = \delta - \alpha \Delta y_t^i + \gamma \Delta Y_t^{us} + \theta Y_{t-1}^{us} + v_t, \tag{6}$$

with the long-run parameters

$$\begin{aligned} \delta &= \beta_0/(1-\beta_3) & \theta &= (\beta_1+\beta_2)/(1-\beta_3) \\ \alpha &= \beta_3/(1-\beta_3) & v_t &= e_t/(1-\beta_3), \\ \gamma &= \beta_1/(1-\beta_3) & & \end{aligned}$$

where Δ is the first difference operator and v_t is an error term. The major drawback of (6) is that it cannot be estimated by OLS, since the first difference of the LHS-variable will be correlated with the error term v_t. Therefore, the appropriate estimation technique is by instrumental variables (IV).

III. Empirical Results

I confine the analysis to a small set of industrialized countries which are large and of comparable size with respect to their population. The reason is that an empirical test of the convergence hypothesis is appropriate only for countries with a similar institutional framework and without geographical peculiarities. Here, it is hoped that particular regional effects may cancel out on average. These countries are Germany, France, Italy, the United Kingdom, and Japan, which are analyzed with respect to catching-up in terms of per capita incomes relative to the United States.

The data for the empirical analysis come from the PWT5 dataset,[11] which provides entries for the period 1950–88. For testing the convergence hypothesis, I use the time series for real GDP per capita in current international prices,[12] which is the appropriate measure for an international comparison of standards of living since it allows for deviations in international purchasing power. For each year, this GDP measure is directly comparable across countries.

The empirical analysis starts with testing whether (3) represents a cointegrating relationship. I use three alternative test procedures to check whether the residual z_t contains a unit root: the augmented

[10] See Kennedy [1992, p. 264] for a hint how to derive (6) from (5).

[11] This set of data is available on personal computer diskettes and through BITNET; all cross-section studies referred to in Section I used this set of data.

[12] Compare column 9 in the PWT5 tables [Summers and Heston, 1991], which is labelled CGDP.

Table 1 – *Testing for Cointegration*

	ADF[a]	Z_α[b]	SP[c]
France	−1.16	−2.33	−1.35
Germany	−2.79	−5.11	−1.17
Italy	−2.37	−5.86	−1.82
Japan	−1.26	−1.97	−1.12
UK	−2.76	−12.92	−2.45

[a] Equation for the augmented Dickey-Fuller test: $\Delta z_t = \alpha_0 z_{t-1} + \alpha_1 \Delta z_{t-1} + e_t$; $H_0 : \alpha_0 = 0$. Critical value [Phillips and Ouliaris, 1990, p. 190]: −2.86 (5%). – [b] Equation for the Phillips-Z_α-test: See Phillips and Ouliaris [1990, p. 171]. Critical value [Phillips and Ouliaris, 1990, p. 189]: −20.49 (5%). – [c] Equation for the Schmidt-Phillips test: See Schmidt and Phillips [1992]. Critical values are available for unit root tests only: approx. −3.15 (5%); critical values are necessarily higher for cointegration tests.

Dickey-Fuller test (ADF) [Said and Dickey, 1984], the Z_α test [Phillips, 1987], and the Schmidt-Phillips test (SP) [Schmidt and Phillips, 1992]. The latter two are less restrictive since they allow for non-i.i.d. errors in the data-generating process of z_t (Z_α test) and for a deterministic misspecification of (3) (SP test). Table 1 contains the results.

All test procedures indicate that the residual z_t of (3) contains a unit root, since the estimated t-ratios are not smaller than the respective critical values. This finding holds true even if the level of statistical significance is reduced from 5 percent to 15 percent. Therefore, the per capita incomes of the US and the other countries seem to follow independent random walks. Put differently, no stable long-run equilibrium relationship between the per capita incomes of these countries seems to exist. Thus, (3) could be considered as representing an entirely spurious regression, pointing to the non-existence of a catching-up process. This result is compatible with the new growth models, but not with the traditional model, at least for the countries in the sample. However, as was noted in the previous section, reasonable parameterizations for the traditional model suggest that the cointegration approach may be inappropriate when applied to the time span since the second world war. Hence, (5) is used as an alternative empirical model for testing the catching-up hypothesis.

This alternative empirical analysis starts with diagnostic checks of (5). I test the possible misspecification of (5) by the Plosser-Schwert-White differencing test (PSW), which needs a minor modification to be applicable for regression equations with lagged dependent variables;[13] and I use the Breusch-Godfrey LM-test[14] (BG) to check for serial correlation in the errors. Table 2 contains the results. The equation for France is rejected by the PSW test. However, this rejection does not necessarily mean that the cointegration approach (equation (3)) represents the relevant empirical model. The rejection may also be due to an implicit higher order dynamic model. Given the relatively small sample size, testing for higher order dynamic models is somewhat restricted. Therefore, the equation for France is not considered for further analysis. Here, it is sufficient to show that a relatively simple dynamic model (AD 1,1) provides a reasonable alternative to the cointegration approach, which uniformly rejected the equations for all countries. That is, the equations for Germany, Italy, Japan, and the UK pass the PSW test, at least at the 1 percent level of statistical significance. Furthermore, all equations pass the BG test at the 1 percent level of statistical significance. Evaluated at the 5 percent level, however, the results point to first-order autocorrelation in the case of Germany and third-order autocorrelation in the case of the UK, but the estimated F-values do not exceed the critical F-values by far. Hence, (5) can be considered as a reasonable alternative to (3), except for the case of France.

Table 2 – *Testing for Misspecification and Autocorrelation*

	Plosser-Schwert-White test[a,c]	Breusch-Godfrey test[b,d]		
		AR(1)	AR(2)	AR(3)
France	6.11	2.70	2.88	1.54
Germany	2.64	4.91	0.44	0.61
Italy	3.72	3.73	0.24	0.13
Japan	1.68	1.82	1.42	1.00
UK	2.43	0.69	0.67	4.43

[a] Test equations (PSW test): $y_t^i = \beta_0 + \beta_1 Y_t^{US} + \beta_2 Y_{t-1}^{US} + \beta_3 y_{t-1}^i + u_t$; $y_t^i = \beta_0^* + \beta_1^* Y_t^{us} + \beta_2^* Y_{t-1}^{us} + \beta_3^* y_{t-1}^i + \beta_4^* y_{t-2}^i + \beta_5^* (Y_{t+1}^{us} + Y_{t-1}^{us}) + \beta_6^* (Y_t^{us} + Y_{t-2}^{us}) + u_t^*$; H_0: $\beta_4^* = \beta_5^* = \beta_6^* = 0$. – [b] Test equation (BG test): $y_t^i = \beta_0 + \beta_1 Y_t^{US} + \beta_2 Y_{t-1}^{US} + \beta_3 y_{t-1}^i + u_t$; $u_t = \beta_0 + \beta_1 Y_t^{us} + \beta_2 Y_{t-1}^{us} + \beta_3 y_{t-1}^i + \sum_{j=1}^{3} \varrho_j u_{t-1} + e_t$; H_0: $\varrho_j = 0$. – [c] Critical values: F (3,29) = 2.93 (5%) and 4.54 (1%). – [d] Critical values: Chi^2 (1) = 3.84 (5%) and 6.63 (1%).

[13] See Maddala [1992] for a textbook exposition.

[14] For a textbook exposition, see, e.g., Johnston [1984] or Maddala [1992].

The next step in the analysis is to check whether (5) actually describes an AD (1,1) model or a serial correlation model of the form

$$y_t^i = c + \theta Y_t^{us} + u_t \quad \text{with} \quad u_t = \varrho u_{t-1} + e_t. \tag{7}$$

Hendry and Mizon [1978] show that this model can be rewritten as

$$y_t^i = (1-\varrho)c + \theta Y_t^{us} - \theta \varrho Y_{t-1}^{us} + \varrho y_{t-1}^i + e_t, \tag{8}$$

which is equivalent to (5) except for the parameters. That is, if the restriction

$$\beta_3 \beta_1 + \beta_2 = 0 \tag{9}$$

holds, then (5) actually describes the serial correlation model of (7). Such a model can be estimated by the Cochrane-Orcutt or the Hildreth-Lu procedure, whereas the AD (1,1) model can be estimated by OLS.

I use the likelihood ratio (LR), the Wald (W), and the Lagrangian multiplier (LM) test[15] to check restriction (9), which discriminates between the models. For linear regression models the LR, W, and LM tests are related in such a way that it is generally possible to reject restriction (9) by the W test but not by the LM test. Table 3 shows, however, that for all countries restriction (7) is rejected even by the LM test at the 5 percent level of statistical significance; restriction (9) is rejected at the 1 percent level of statistical significance by the W test. Therefore, the data can be adequately described by an AD (1,1) model, not by a serial correlation model; point estimates of the long-run parameters may be derived from an IV-estimation of (6).

Obviously, the results of an IV-estimation critically depend on the properties of the selected instruments. For instance, a low or a negative R^2 from an IV-regression indicates that something is wrong with the specification of the model or with the selection of the instrument. Therefore, I use two different instruments to estimate (6) in order to check the robustness of the results. The upper part of Table 4 contains the resulting parameter estimates when ΔY_{t-1}^{US} is chosen as an instrument for Δy_t^i. Apparently, this is not a good choice for the UK equation. The lower part of Table 4 contains the parameter estimates when the sum of the differenced LHS-variables absent from the equation under consideration ($\sum_{i \neq j} \Delta y_t^j$) is chosen as an instrument for Δy_t^i. This instrument yields a significant R^2 for the UK equation, but otherwise lower R^2s except for the case of Italy. The results for

[15] For a textbook exposition, see, e.g., Maddala [1992].

Table 3 – *Testing Serial Correlation vs. Misspecified Dynamics*

	LR	*W*	*LM*
Germany	6.42	6.99	5.91
Italy	6.57	7.17	6.03
Japan	7.52	8.32	6.83
UK	18.24	23.41	14.49

Note: The test equations are the following:

$$LR = n \log_e \left(\frac{RRSS}{URSS}\right), \qquad W = n\frac{RRSS-URSS}{URSS}, \qquad LM = n\frac{RRSS-URSS}{RRSS},$$

where n is the number of observations, RRSS is the sum of squared residuals from equation 5 (estimated by Cochrane-Orcutt), and URSS is the sum of squared residuals from equation 3 (estimated by OLS). – Critical value: $Chi^2(1) = 3.84$ (5%).

Table 4 – *Point Estimates for the Long-Run Parameters*[a]

	δ	θ	R^2
		1. IV: ΔY_{t-1}^{US}	
Germany	−0.044 (0.203)	0.082 (0.019)	0.581
Italy	−0.545 (0.084)	0.125 (0.010)	0.852
Japan	−1.431 (0.168)	0.207 (0.015)	0.917
UK	0.291 (0.098)	0.046 (0.016)	−0.878
		2. IV: $\sum_{i \neq j} \Delta y_t^j$	
France	−0.375 (0.106)	0.112 (0.012)	0.745
Germany	−0.398 (0.164)	0.119 (0.018)	0.552
Italy	−0.633 (0.079)	0.135 (0.009)	0.863
Japan	−1.611 (0.131)	0.230 (0.015)	0.885
UK	0.354 (0.034)	0.032 (0.004)	0.618

[a] Standard errors in parentheses.

Germany should be interpreted cautiously, because of the relatively low R^2.

Turning to the long-run parameter estimates, one finds that all countries are catching up to the US, since θ is positive in all equations. With this result the non-cointegration finding of Table 1 may be reinterpreted as the acceptance of a possibly false hypothesis. For instance, testing for cointegration by an analysis of the residual z_t of the static model in (3) may involve a relatively high probability of committing a type II error when the time span under consideration actually reflects off-steady-state behavior. Then, it will be impossible to statistically discriminate between the hypothesis of a non-stationary residual (no cointegration) and a serially correlated residual (wrong functional form, misspecified dynamics, or serial correlation model).

A unit root in the residuals and a high degree of autocorrelation are observationally equivalent for reasonable sample sizes. While the former is consistent with the new growth models, the latter is inconsistent with the traditional growth model only if this model predicts a high speed of convergence to the steady state path after an exogenous shock. However, theoretical considerations and empirical results based on cross-section studies [16] point to a relatively slow rate of convergence: A fair estimate is that an average economy will reach halfway to steady state in about 35 years. Thus, the data used in this paper may mainly reflect off-steady-state behavior. For an empirical analysis of this time span, the implication is to begin with a general dynamic model, and not to give too much weight to the results of cointegration tests, which are valid for large samples only. Hence, the failure to find a cointegrating relationship between the per capita incomes of the US and other developed countries does not necessarily support the new growth theories.

The estimates for the regression constant (δ), also presented in Table 4, can be interpreted in terms of the steady state levels of per capita incomes. The statistically significant negative constants for Italy and Japan indicate a higer steady state level of per capita income in these countries relative to the US, and the positive constant for the UK indicates a steady state level of per capita income below that of the US. The results for Germany depend on the instrument being chosen; a statistically insignificant constant indicates a convergence to the US level of per capita income. Taken together, these results confirm the hypothesis of conditional convergence. They do not necessarily imply, however, a falsification of the new growth models.

IV. Conclusion

The basic message of traditional constant-returns-to-scale models of economic growth is that market forces will ensure a catching-up of

[16] See Footnote 6 in Section II.

per capita incomes between rich and poor countries, given that the countries under consideration do not differ too much with respect to their institutional arrangements and time preferences. This message is not necessarily confirmed by the "new" growth models. They can explain why international differences in per capita incomes may persist, even if the countries under consideration have access to the same technology, and the international mobility of capital is not restricted. Thus, the new models predict that market forces alone might not be sufficient to initiate a catching-up process of poor countries.

The time-series evidence based on the newly introduced concept of cointegration analysis seems to support the new models. However, these results are based on a very restrictive econometric framework. Less restrictive model specifications and estimation techniques used in this paper produce results that are more favorable for the traditional model. The catching-up hypothesis cannot be rejected for a number of countries when the econometric model allows for conditional convergence of per capita incomes over time, due to the selection of an appropriate functional form and an explicit modeling of dynamic adjustment processes. This finding shows that the application of an inappropriate econometric approach may easily lead to the acceptance of a probably false hypothesis. Therefore, the empirical evidence does not necessarily support the recommendation of interventionist economic policies to achieve a catching-up process, which is tempting to be derived from the new growth models. But obviously it does not necessarily support the alternative hypothesis as well. The time-series evidence is shown to be as inconclusive as the cross-section evidence. Hence, attempts to empirically discriminate between alternative growth models seem to lead to dead ends.

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A b s t r a c t: Empirical Evidence for Alternative Growth Models: Time Series Results. – Recent attempts to discriminate between alternative models of economic growth have led to a full circle of results by focussing on cross-section data. In this paper, the author shows that the time series evidence is inconclusive as well, depending on the specification of the test equation with respect to its functional form and its dynamic modelling, and on the estimation technique used. Hence, given the data at hand, attempts to devise explicit tests of alternative growth models seem to lead to dead ends.

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Z u s a m m e n f a s s u n g: Empirische Befunde für alternative Wachstumsmodelle. Ergebnisse aus Zeitreihenanalysen. – Jüngere Versuche, mit Hilfe von Querschnittsdaten zwischen alternativen Wachstumsmodellen zu unterscheiden, haben zu widersprüchlichen Ergebnissen geführt. In diesem Aufsatz zeigt der Verfasser, daß die Befunde aus Zeitreihenanalysen ebenfalls nicht schlüssig sind und davon abhängen, wie die Testgleichung im Hinblick auf ihre funktionale Form und ihre dynamische Gestalt spezifiziert wird und welche Schätztechnik benutzt wird. Bei den gegebenen Daten scheinen deshalb Versuche, Tests für alternative Wachstumsmodelle explizit zu entwerfen, in die Irre zu führen.