MODELING GROWTH (AND LIBERALIZATION) USING SMOOTH TRANSITIONS ANALYSIS

DAVID GREENAWAY, STEPHEN LEYBOURNE, and DAVID SAPSFORD*

Economic liberalization has been a pervasive phenomenon over the last twenty years. Programs have been initiated on the assumption that liberalization promotes economic growth, but the empirical evidence for this is limited. This paper takes a novel approach to modeling growth and structural change as smooth transitions. This allows us to model deterministic change without imposing discrete changes. We use smooth transition analysis to reappraise the time-series properties of long-run growth rates in a number of developing countries which have undertaken liberalization. Our results challenge conventional wisdom on both methodological and empirical grounds. (*JEL* F1, C2, O4)

I. INTRODUCTION

There is a large and growing literature on liberalization and its effects, stimulated initially by the very extensive trade reform programs promoted in developing countries over the last fifteen years or so and sustained by the attempts to understand what is happening in the eastern and central European economies. The latter do not feature in this paper, though the techniques demonstrated here may in time prove to be useful for gaining insight into the effects of regime changes underway.

Liberalization, in the sense of trade reforms which reduce anti-export bias in one way or another while improving the information content of relative price changes, has been promoted in some eighty or so developing countries in the 1980s and 1990s. Some of these liberalizations are unilateral, most are policy conditioned under the aegis of World Bank Structural Adjustment Loans (SALs).[1] All have been undertaken on the assumption that liberalization will ultimately improve export and growth performance. The evidence on liberalization and growth, however, is at best somewhat mixed. On the one hand, one has the exhaustive study of Papageorgiou, Michaely and Choksi [1991] which claims a close and direct association. On the other hand, Greenaway and Sapsford [1994] find only a limited role for liberalization, and a range of studies of SALs associate adjustment with a deterioration in growth performance.[2]

These studies generally rely either on crude data inspection, simple correlation analysis, or multivariate regression. This paper takes a new approach to the question. It starts from the presumption that any changes in economic performance following a liberalization may be more appropriately modeled as a steady transition rather than a discrete change. To model

ABBREVIATIONS

PMC: Papageorgiou, Michaely and Choksi [1991]
LSTR: Logistic smooth transition regression
DFR: Dickey-Fuller regression

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1. For a review of experience, see Whalley [1991] and Greenaway and Morrissey [1993].

2. See Harrigan and Mosley [1991] and World Bank [1990].

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the change in this way, we make use of the recent work of Granger and Terasvirta [1994], who explore the properties of a variety of non-linear specifications that facilitate modeling structural change as a smooth transition between states. We estimate and test the adequacy of a number of such specifications.

Our analysis is in two stages. First we take a novel approach to the modeling of growth which allows us to model deterministic change without, as other analysts have done, imposing discrete changes. Not only does this help clarify the statistical properties of the time series, it also challenges the assumptions that underpin much growth modeling. Having identified the transitions in the growth series we are investigating, we then explore the coincidence of these transitions with well-documented episodes of liberalization. We do not formally test whether liberalization results in growth. Our results are, however, informative in two respects. Firstly, they place a question mark against the widely held presumption that liberalization is a panacea for growth. Secondly, they also point the way to the appropriate econometric modeling of these processes.

The remainder of the paper is structured as follows. Section II briefly describes the context against which our analysis is set. Section III sets out the details of our methodology and the models to be tested. Section IV reports our results, contrasts them with those of previous work and evaluates the implications for policy. Section V concludes and identifies possible extensions.

II. LIBERALIZATION, (EXPORTS) AND GROWTH: PREVIOUS EVIDENCE

There are essentially two strands of the literature which provide relevant background: one which relates exports and growth; a second which relates liberalization to exports and growth.

The exports and growth literature is an extensive one, which goes back some years.[3] It starts from the presumption that exports and growth are directly related, with causality running from changes in exports to changes in growth.[4] Analysts have deployed a variety of empirical methods, usually of growth accounting type models, and covering a range of countries and time periods. The consensus from this literature can be summarized as follows. First, it seems to be the case that exports and growth are in general correlated. However, the correlation holds rather more strongly for cross-section than time-series data. Second, there is substantial evidence in favor of a threshold effect, i.e. industrialization needs to have proceeded beyond some critical level, (as proxied by GDP per capita), for the export-growth relationship to hold. Thus notwithstanding the qualifications relating to causality which some authors have raised, there is quite a lot of empirical support for the assertion that exports and growth are related. This has been taken by some as a basis for recommending liberalization, the presumption being that liberalization stimulates export performance and this in turn stimulates growth performance.

More recent are attempts to establish directly the liberalization-exports-growth relationship. In policy terms the most influential study here is Papageorgiou, Michaely and Choksi [1991] (hereafter PMC). This is a massive study of 36 liberalization episodes in 19 countries, over the period from the mid 1950s to the mid 1970s. Using essentially informal analysis which compares average export and growth performance across all of the episodes for the three years before and three years after each episode, they conclude, quite unequivocally, that liberalization boosts both exports and growth. Notwithstanding the reservations about the methodological foundation to this work articulated in Greenaway [1993] and Collier [1993], the results have been widely reported as conclusive proof of the efficacy of liberalization—most notably of course in the World Bank itself.[5]

Greenaway and Sapsford [1994] subject PMC sample (or rather a subset thereof), to more rigorous statistical scrutiny using slope and intercept dummies within a standard pro-

3. A recent survey of results is given in Greenaway and Sapsford [1994].

4. Note that some studies have conducted formal tests and question whether causality unambiguously runs from exports to growth. See, for instance, Jung and Marshall [1985] and Darrat [1987].

5. See, for example, Thomas and Nash [1992].

duction function growth model.[6] The model was estimated on data for 13 of the 19 countries in the PMC sample.[7] The results indicate that in some two-thirds of all cases, liberalization appears to have no discernable impact on the exports-growth relationship. In three out of the 12 cases, a significant positive relationship is found, in one case a significant negative relationship. Given the conviction with which the PMC results are reported, and the enthusiasm with which they have been embraced by the key lending agency, these results are unsettling. It could of course be that liberalization does not affect growth. A second possibility is that the results are an artifact of a misspecified model and, given the reservations of Levine and Renelt [1992] on most growth modeling, this is quite plausible. Another possibility is that liberalization does not affect economic performance via a discrete break, but rather by initiating a transition. If so, structural break tests which rely on the existence of a discrete break will generally fail to find one. Most of the case study evidence suggests that trade reforms initiate transitions rather than discrete breaks and we need therefore to model the process as such, ideally without imposing priors on the data. Smooth transitions analysis offers that facility.

III. SMOOTH TRANSITIONS ANALYSIS

Smooth transition analysis is an approach to modeling deterministic structural change in a time-series regression. Originally proposed by Bacon and Watts [1971] and Maddala [1977, 396], it has been more recently developed by Lin and Terasvirta [1994] and Granger and Terasvirta [1993, ch. 7]. The basic idea is quite simple. Rather than attempt to identify any change as a single structural break, one identifies it as a smooth transition between regression regimes over time.

Where liberalization is concerned, this is intuitively appealing. Even where liberalization is implemented in a "big bang," any subsequent effects on growth will typically be gradual rather than automatic, the speed of adjustment being dependent on the efficiency of markets in the economy in question. In developing countries, big bang liberalizations are the exception rather than the rule; sequenced liberalizations are more common. In these circumstances it can take time for the reforms to gain credibility and for agents to react. Either way, modeling the impact on growth as a discrete structural break is inappropriate.

More concretely, following Granger and Terasvirta [1993, ch. 7], a simple logistic smooth transition regression (LSTR) trend model may be written as:

$$(1) \qquad \ln(y_t) = \alpha_1 + \alpha_2 S_t(\gamma, \tau) + \beta_1 t + \beta_2 t S_t(\gamma, \tau) + \varepsilon_t \,, \; t = 1, \ldots, T$$

$$S_t(\gamma, \tau) = \{1 + \exp[-\gamma(t - \tau)]\}^{-1}.$$

Here S_t is the well-known curvilinear logistic function that maps t onto the interval (0, 1) and ε_t is a zero mean disturbance term. Under this formulation and assuming $\gamma > 0$, the model transition occurs smoothly between the initial state

$$\ln(y_t) = \alpha_1 + \beta_1 t + \varepsilon_t \,, \; t \to -\infty$$

and the final state

$$\ln(y_t) = (\alpha_1 + \alpha_2) + (\beta_1 + \beta_2) t + \varepsilon_t \,, \; t \to +\infty$$

corresponding to $S_{-\infty} = 0$ and $S_{+\infty} = 1$, respectively. Hence, the mean growth rate of y_t, which is the coefficient on the trend variable t, changes from β_1 to $(\beta_1 + \beta_2)$ through time. Notice the model simultaneously allows the intercept to change from α_1 to $(\alpha_1 + \alpha_2)$. Here, τ is a location parameter which determines the timing of the transition. For $t = \tau$, we have

$$\ln(y_t) = (\alpha_1 + 0.5\alpha_2) + (\beta_1 + 0.5\beta_2)t + \varepsilon_t$$

such that τ identifies the transition midpoint. The velocity of transition is controlled by the parameter γ. If γ takes a large value then the

6. Both intercept and slope dummies were used because liberalization has the potential to impact on both the level and rate of growth of GDP. The former is typically association with the once and for all benefits of improved resource allocation; the latter is attributable to more rapid factor productivity growth in export-oriented sectors.

7. Brazil, Columbia, Greece, Israel, Korea, New Zealand, Pakistan, Peru, Philippines, Spain, Sri Lanka, Turkey, Yugoslavia.

transition is completed in a short period of time and as γ tends to infinity the model collapses to one with an instantaneous structural break in intercept and trend at time $t = \tau$. Thus (1) embeds the standard structural break model as a special case. The parameters α_2 and β_2 determine the direction of transition in the intercept and trend, respectively. If $\gamma < 0$, the initial and final model states are reversed but the interpretation of the parameters remains the same.

The model (1) is nonlinear in parameters and may be estimated by nonlinear least squares (NLS) using a suitable iterative optimization algorithm. As pointed out in Granger and Terasvirta [1993, ch. 7], while the other parameter estimates can converge quickly, that for γ may converge very slowly, particularly if the true parameter value is large (such that the transition occurs quickly). This is because a large set of estimated values of γ lead to very similar values of S_t, which deviate noticeably from each other only in a local neighborhood of the location parameter τ. The practical consequence of this is that standard errors of the NLS estimate of γ may appear artificially large and should not, therefore, be taken necessarily to indicate insignificance of the estimate.

The logistic function S_t as specified here does impose certain restrictions, in that the transition path is monotonic and symmetric around its midpoint. More flexible specifications could also be considered which, for example, could allow for non-monotonic and non-symmetric transition paths. This is facilitated by including a higher order polynomial in t in the exponential term of S_t. In addition, we constrain the transitions in intercept and trend to occur once only, simultaneously, and with the same velocity. Clearly, a specification which does not impose these restrictions could also be entertained. However, a particular advantage our specification has over more complex specifications is that all the parameters have very straightforward interpretations; in more heavily parameterized versions this is no longer the case. Moreover, since the number of observations available in this study is relatively small, degrees of freedom problems would also quickly arise. For these reasons we do not attempt any such extensions here.

In sharp contrast to conventional approaches to modeling structural change, no a priori information is used to fix the date of a transition since the midpoint of the transition is determined endogenously via the parameter τ (with the parameter γ then effectively identifying the start and end points). From the standpoint of modeling liberalization episodes, what this means is that the data are allowed to determine all the pertinent features of any transition in the real growth rate—its timing, duration and direction. If any such transition is found, and it need not be, one can then refer back to the dating of a liberalization episode, as established from policy accounts, to see whether or not there is any apparent coincidence of timing. Our central focus is the PMC study where a number of distinct episodes are identified by the authors. The specific questions we are interested in are: Is there any evidence of a transition in growth rate in the countries in question over the period to which the PMC study applies? Is there any connection in terms of timing between the transition and liberalization as identified by PMC?

IV. MODEL ESTIMATION

The LSTR model was estimated using annual time series data for 13 countries taken from the PMC sample: Brazil, Colombia, Greece, Korea, Israel, New Zealand, Pakistan, Portugal, Spain, Sri Lanka, Argentina, Yugoslavia and Indonesia. The dependent variable y_t was real GDP per capita at 1980 purchasing power parity prices. To account for the possibility of stochastic dynamics, the estimated model was augmented to include a lagged dependent variable term as

$$\ln(y_t) = \alpha_1 + \alpha_2 S_t(\gamma, \tau) + \beta_1 t + \beta_2 t S_t(\gamma, \tau) + \varphi \ln(y_{t-1}) + \varepsilon_t. \tag{2}$$

Including (at most) a single lagged term in $\ln(y_t)$ was found to be sufficient to yield serially uncorrelated residuals for each series. Notice that with this augmented version of the model, the mean growth rates of y_t corresponding to the initial and final model states are now given by $\beta_1/(1-\varphi)$ and $(\beta_1+\beta_2)/(1-\varphi)$, respectively.

The Berndt-Hall-Hall-Hausman optimization algorithm in GAUSSX 3.2 was used to compute the NLS estimates of the seven unknown parameters in model (2) for each country. Since the model is linear in the parameters α_1, α_2, β_1, β_2, and φ, considerable economy in estimation is possible as these can be "concentrated out" of the sum of squares function using OLS. The estimation results are given in Table I under LSTR. Asymptotic t-ratios for the parameter estimates are given in parentheses. Where particular parameter estimates are not present, they were found to be insignificantly different to zero (at the 5% level), and the results therefore refer to the model re-estimated assuming these parameters are equal to zero. We also estimate model (2) imposing the restriction $S_t = 0$ such that the intercept and trend coefficients are constant over time (no transition between regimes occurs). Given the inclusion of the lagged dependent variable term, (2) then specializes to the well-known Dickey-Fuller regression (DFR), which allows the possibility of testing for a unit autoregressive root, or stochastic trend, in real GDP ($\ln y_t$) against the alternative of stationarity around a linear deterministic trend. We report these regression results under DFR.

The standard likelihood ratio test for the restriction $\gamma = 0$ does not provide us with a valid test of the null hypothesis of constancy of the intercept and trend against the smooth transition alternative. This is because under this null the parameters α_2, β_2 and τ are no longer identified (i.e., they may assume any possible value). However, a valid Lagrange multiplier test of this hypothesis has been suggested by Lin and Terasvirta [1994]. This test is based on a two-step approach proposed by Davies [1977]. Briefly, the test procedure first assumes that the logistic function S_t can be adequately approximated by a polynomial function of t up to some order k, say, via a Taylor series expansion. Next, the residuals from the DFR model (which assumes constant intercept and trend) are constructed, together with the residual sum of squares which we denote SSR_0. These residuals are then regressed on the same DFR regressors together with additional regressors which are polynomial terms in t up to order $k+1$. Denoting the sum of squared residuals from this second regression as SSR_1, the Lagrange multiplier test has the form

$$LM = (SSR_0 - SSR_1) / (SSR_0 / T).$$

Given standard regularity conditions LM has an asymptotic $\chi^2(k)$ distribution under the null hypothesis of constancy of the intercept and trend. The degrees of freedom of the limiting distribution is k and not $k+1$ because t itself appears as a regressor in the null DFR model.

For our purposes, we assume a third-order Taylor-series expansion of $S_t(\gamma, \tau)$ is adequate, requiring that polynomial terms in t up to the fourth order are included in the second-stage regression. The reported LM statistic therefore has a $\chi^2(3)$ distribution under the null hypothesis of constancy. Based on the 5% significance level of a $\chi^2(3)$ distribution (critical value 7.81), the results of the LM tests in Table I suggest that for all 13 countries evidence of a transition in intercept or trend (or both) is present. Looking at the estimated models more closely, we see that only for Brazil is the transition restricted to the intercept term; all the other countries display evidence of a transition in trend, and hence in mean growth rates (often in addition to a transition in the intercept term).

For each estimated model we also report the value of the Box-Pierce $Q(3)$ statistic for residual autocorrelation. Under the null of zero autocorrelation this statistic has an approximate $\chi^2(2)$ distribution (5% critical value 5.99) if the estimated model contains a lagged dependent variable, and a $\chi^2(3)$ distribution if it does not. At the 5% level, the null of zero residual autocorrelation is not rejected for any of the fitted LSTR models. This is in spite of the fact that eight out of the 13 models actually contain no lagged dependent variable term.

Under DF we give the value of the Dickey-Fuller unit root test (i.e., the t-test that $\varphi = 1$) for all the models where φ is not constrained to be zero. In the case of the LSTR, the nonlinearity in the deterministic components means this cannot be considered as a formal test for a unit root, since the t-ratio will no longer have the standard Dickey-Fuller distribution. It is included, nonetheless, for the purpose of comparison.

In the DFR models the null hypothesis of a unit root in $\ln(y_t)$ cannot be rejected at the

5% level for any of the countries in the sample with the exception of Indonesia (an approximate 5% critical value for DF is –3.55). In the five LSTR models where φ is estimated, as a consequence of the estimates of φ being much closer to zero than in the corresponding DFR models, the values of the DF are now considerably lower (and less than the critical value –3.55 in four of the five cases). The implication is that there is little support for the unit-root hypothesis in these series.

In terms of the dynamics of these GDP series, therefore, it appears that once deterministic dynamics in the form of smooth transition functions are permitted, then stochastic dynamics either play a greatly reduced role or, as for eight out of the 13 series, no role at all. Moreover, according to the DFR regressions, 12 out of the 13 series would conventionally be characterized as unit-root processes with the implication that GDP has no tendency to revert to a linear mean function. In contrast, our results from the LSTR models show that the movements in real GDP are rather better characterized as stationary fluctuations around a mean path which itself exhibits a smooth nonlinear transition between two distinct linear functions.

As regards GDP growth, the implication is that rather than the mean growth rates being constant over the sample period, as the DFR models can be shown to imply once the unit coefficient on φ is imposed (the growth rate then being given by $\Delta\ln(y_t)$), instead they evolve in a smooth nonlinear fashion between two different values.

A closer examination of the signs of the estimated values of β_1, β_2 and γ shows that the direction of transition of the mean growth rate of real GDP is positive for Colombia, Korea, Sri Lanka and Indonesia. For the remaining countries in the sample (Greece, Israel, New Zealand, Pakistan, Portugal, Spain, Argentina and Yugoslavia) the transition is a negative one. To see these differences in experience more clearly, the estimated mean growth transitions are plotted in Figures 1 to 12 (no plot is included for Brazil as the mean growth rate was found to be constant). The quantity plotted against time is the estimated counterpart of $(\beta_1, + \beta_2 S_t)/(1 - \varphi)$, and hence represents the estimated per capita mean growth rate at each point in time (plots of the actual and fitted values of $\ln(y_t)$ are available from the authors on request; the fitted models appear to capture the main features of the data very successfully). In each case the estimated trend coefficient can be observed to evolve smoothly between limiting values. Also, considerable diversity in the speed and timing of the transitions is evident. For instance, Yugoslavia undergoes the most rapid transition between mean growth paths (γ is estimated to be 3.58) which is more or less fully completed within the period 1976–1980, whereas the Korean transition, centered around 1968, appears to be the most gradual (γ is estimated to be 0.23); in fact, it is not fully completed even over the entire sample period. Importantly, however, none of the transitions appear to occur in an instantaneous fashion (which would be associated with a very large estimated value of γ). This, therefore, raises rather serious doubts about the ability of models which only permit discrete structural breaks, as typically employed in other studies, to capture the salient features of GDP growth in this sample of countries.

Thus, our findings that the GDP series are better characterized as being stationary around a deterministic path which exhibits a smooth nonlinear transition between two linear trend functions challenges not only the assertions of Nelson and Plosser [1982] that these kinds of series typically need to be modeled as unit-root processes (when the alternative process is stationary around fixed linear deterministics), but also those of Perron [1989] that such series can be adequately modeled as being stationary simply by introducing a discrete structural break into a trend function. In addition, our results raise a further important issue with regard to the structural modeling of economic growth: if the dependent variable is found to exhibit characteristics associated with a smooth transition processes, then this has important economic and statistical implications for building structural models of economic growth since any proposed structural model must be fully capable of explaining this kind of behavior.

Finally, to gain some indication as to whether a positive or negative transition in real GDP growth may have been triggered by a liberalization episode, we have imposed PMC episodes as shaded areas over the estimated transition paths. Note that this does not

TABLE I
Estimates from Logistic Smooth Transition Models

		α_1	α_2	β_1	β_2	γ	τ	φ	R^2	LM	$Q(3)$	DF
Brazil	LSTR	4.277	–0.308	0.042		0.889	19.11	0.389	0.997	16.2	1.05	–5.50
(1963–1985)		(5.53)	(–5.26)	(6.08)		(3.77)	(48.1)	(3.52)				
	DFR	0.383		0.001				0.953	0.981		5.54	–0.40
		(0.47)		(0.17)				(8.29)				
Colombia	LSTR	7.158		0.026	–0.017	–0.495	15.72		0.994	10.1	5.40	
(1955–1982)		(390.)		(35.7)	(–6.48)	(–4.53)	(26.9)					
	DFR	1.623		0.008				0.769	0.990		4.86	–3.10
		(3.10)		(3.63)				(10.3)				
Greece	LSTR	3.249		0.028	–0.007	0.642	28.60	0.530	0.995	14.5	0.58	–2.93
(1952–1985)		(2.99)		(2.93)	(–2.65)	(1.87)	(27.5)	(3.31)				
	DFR	0.140		–0.001				0.991	0.992		1.38	–0.12
		(0.25)		(–0.27)				(12.0)				
Korea	LSTR	6.565		0.055	–0.072	–0.230	12.56		0.997	9.66	6.22	
(1955–1979)		(157.)		(25.6)	(–4.09)	(–5.02)	(8.75)					
	DFR	0.595		0.010				0.903	0.995		1.63	–1.71
		(1.71)		(2.87)				(16.0)				
Israel	LSTR	5.170	0.660	0.031	–0.028	1.121	18.45	0.321	0.997	9.92	1.97	–4.25
(1953–1985)		(4.31)	(3.90)	(3.66)	(–3.74)	(1.88)	(26.0)	(2.01)				
	DFR	0.284		–0.001				0.973	0.991		2.78	–0.42
		(0.58)		(–0.49)				(15.2)				
New Zealand	LSTR	4.746		0.014	–0.003	1.784	26.08	0.431	0.989	7.85	0.16	–3.57
(1951–1984)		(3.47)		(3.32)	(–2.99)	(1.06)	(42.5)	(2.61)				
	DFR	1.304		0.002				0.848	0.980		030	–1.65
		(1.70)		(1.26)				(9.20)				
Pakistan	LSTR	6.176		0.071	–0.037	0.681	10.06		0.988	11.6	4.68	
(1960–1984)		(190.)		(8.39)	(–5.22)	(4.50)	(21.7)					
	DFR	1.383		0.005				0.788	0.963		1.36	–1.77
		(1.82)		(1.44)				(6.61)				

TABLE I continued
Estimates from Logistic Smooth Transition Models

		α_1	α_2	β_1	β_2	γ	τ	φ	R^2	*LM*	*Q*(3)	*DF*
Portugal	LSTR	4.272	0.590	0.022	–0.015	0.439	20.01	0.381	0.997	14.2	3.81	–3.65
(1950–1985)		(3.66)	(3.01)	(2.88)	(–2.57)	(2.88)	(18.6)	(2.22)				
	DFR	0.098		–0.001				0.995	0.993		1.76	–0.06
		(0.17)		(–0.24)				(11.3)				
Spain	LSTR	7.480	0.953	0.031	–0.022	0.267	19.97		0.995	14.4	4.00	
(1950–1985)		(265.)	(3.30)	(3.20)	(–3.03)	(2.57)	(11.0)					
	DFR	1.133		0.005				0.856	0.990		0.18	–1.43
		(1.52)		(1.11)				(8.49)				
Sri Lanka	LSTR	6.699	–9.45	0.012	0.031	0.857	24.23		0.954	15.9	0.34	
(1950–1985)		(345.)	(–6.63)	(6.79)	(6.95)	(1.83)	(28.4)					
	DFR	1.123		0.003				0.832	0.899		1.78	–1.16
		(1.17)		(1.53)				(5.75)				
Argentina	LSTR	7.954		0.026	–0.020	0.478	22.27		0.912	13.3	7.27	
(1959–1985)		(298.)		(8.50)	(–4.15)	(2.45)	(25.4)					
	DFR	0.780		–0.002				0.910	0.832		3.62	–0.82
		(0.88)		(–1.03)				(8.29)				
Yugoslavia	LSTR	7.425	0.742	0.049	–0.036	3.575	18.94		0.996	11.7	1.89	
(1960–1985)		(519.)	(6.58)	(36.1)	(–7.13)	(4.20)	(40.0)					
	DFR	1.402		0.007				0.819	0.990		0.63	–1.06
		(1.11)		(0.89)				(4.79)				
Indonesia	LSTR	6.209		0.041	–0.077	–0.351	7.519		0.997	14.1	4.82	
(1962–1985)		(99.5)		(16.5)	(–2.30)	(–4.65)	(4.38)					
	DFR	1.967		0.019				0.668	0.995		3.30	–4.04
		(4.07)		(4.41)				(8.15)				

FIGURE 1
Estimated Growth: Colombia

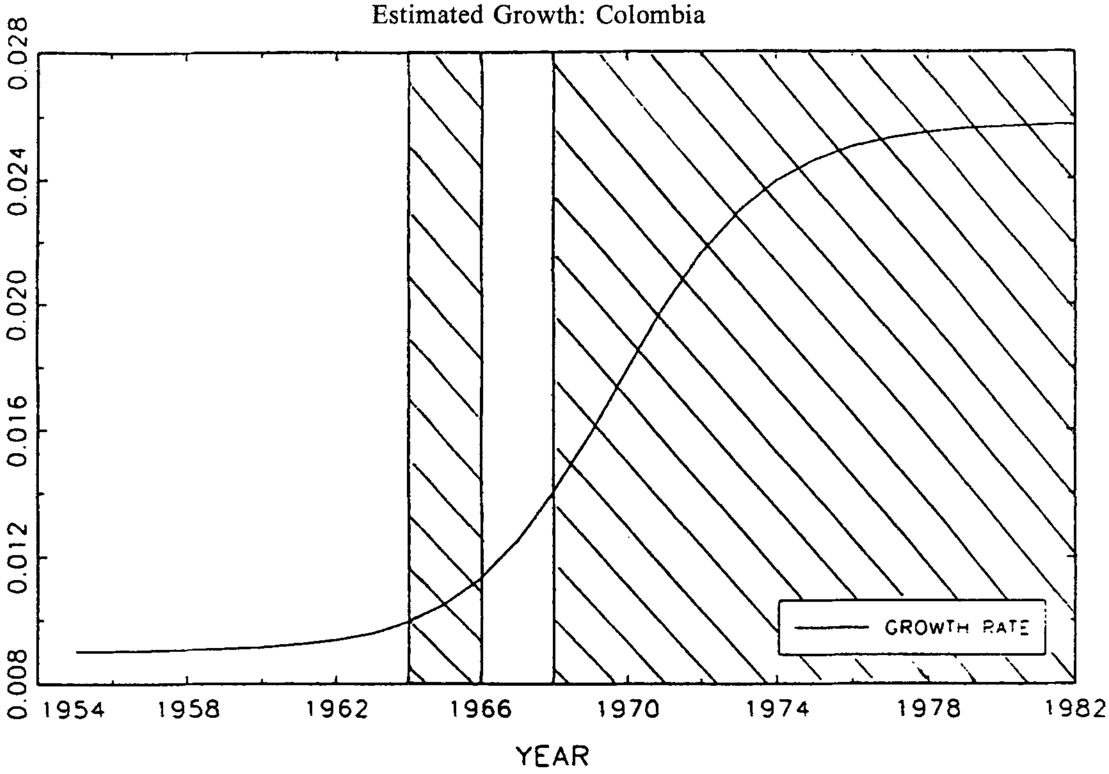


FIGURE 2
Estimated Growth: Greece

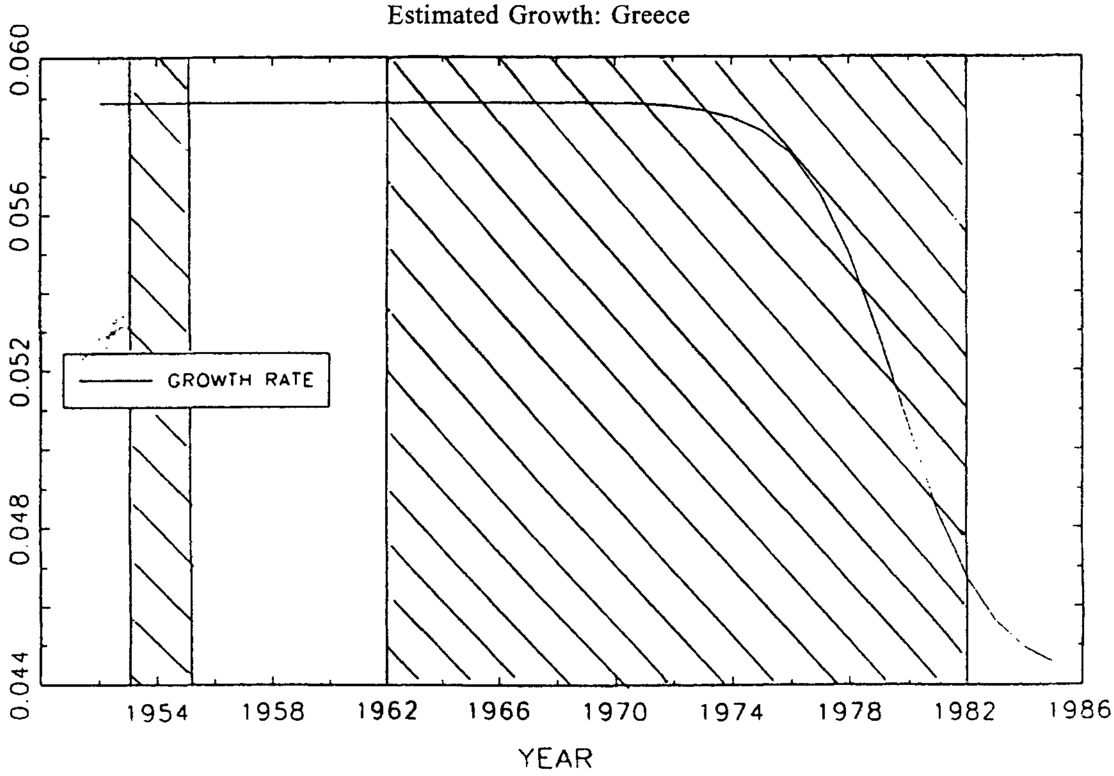


FIGURE 3
Estimated Growth: Korea

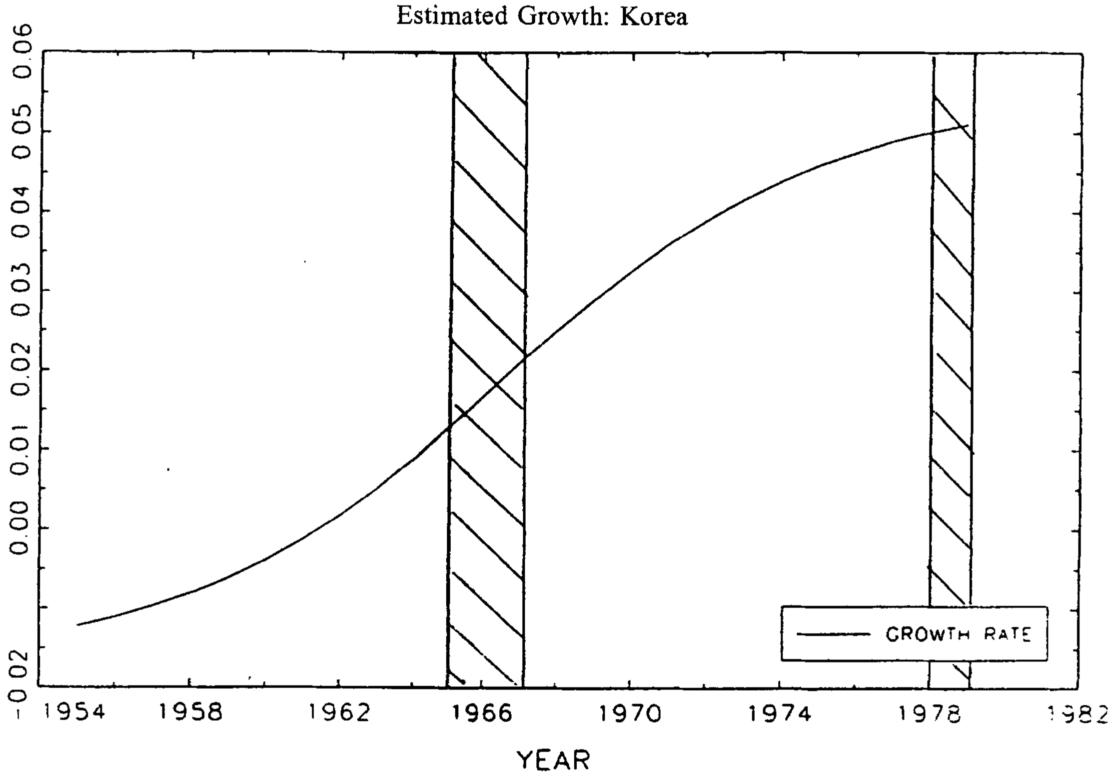


FIGURE 4
Estimated Growth: Israel

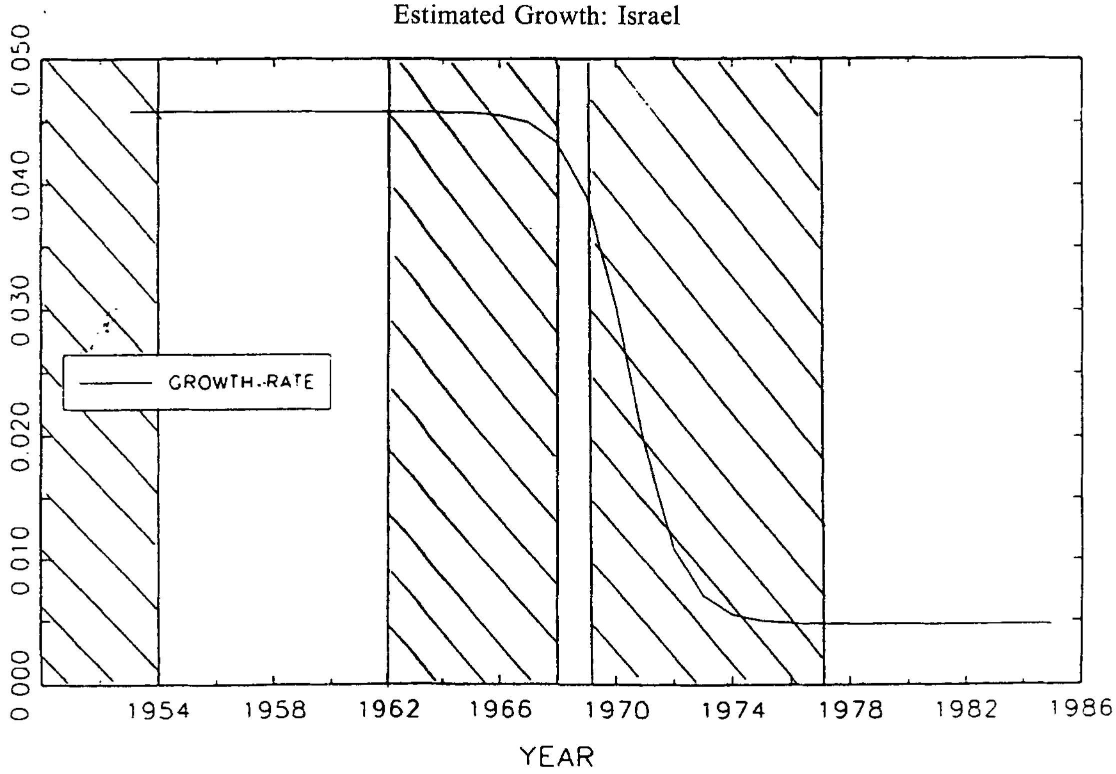


FIGURE 5
Estimated Growth: New Zealand

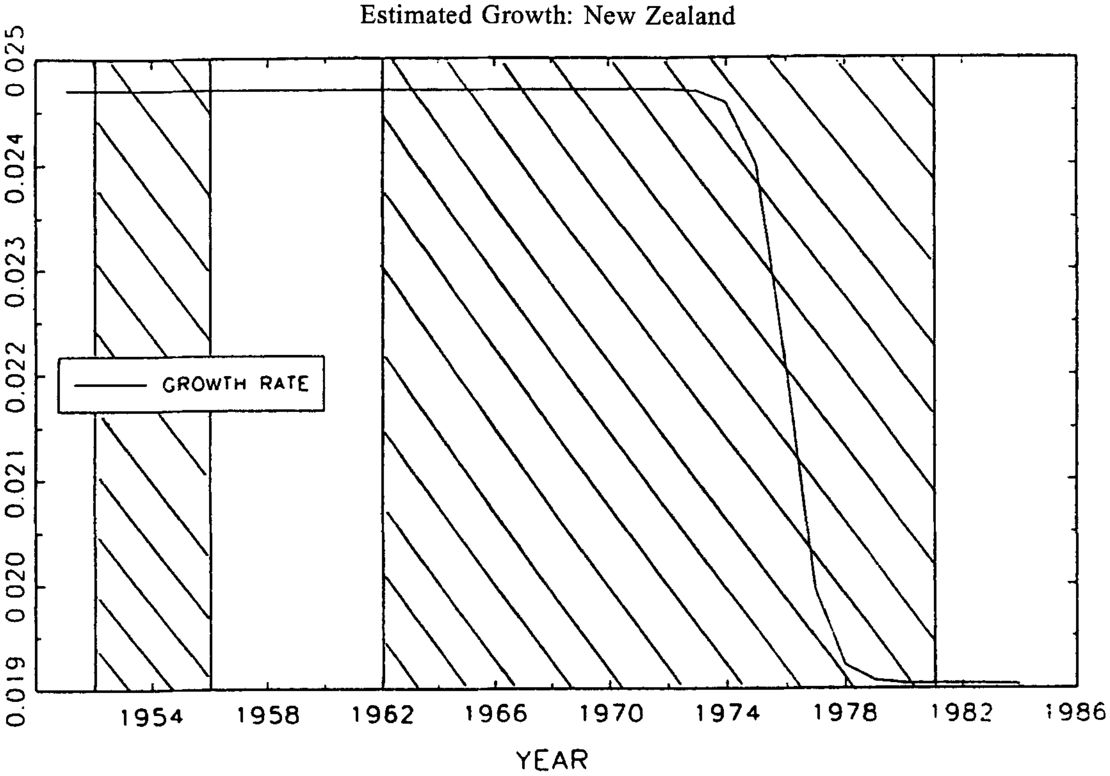


FIGURE 6
Estimated Growth: Pakistan

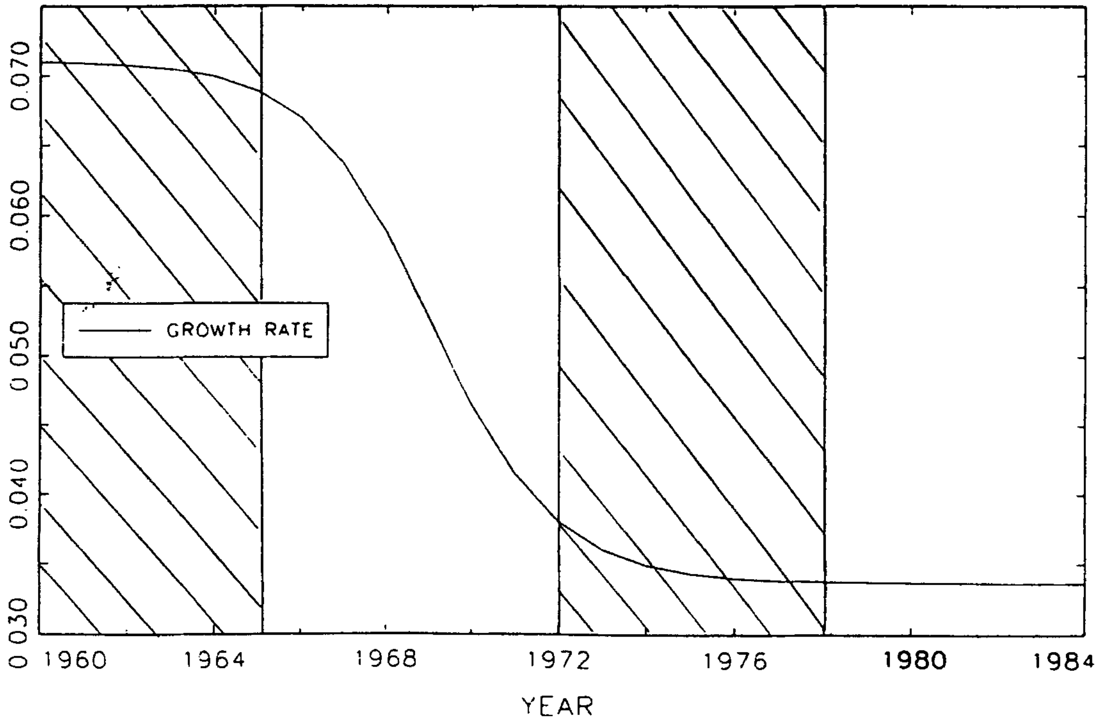


FIGURE 7
Estimated Growth: Portugal

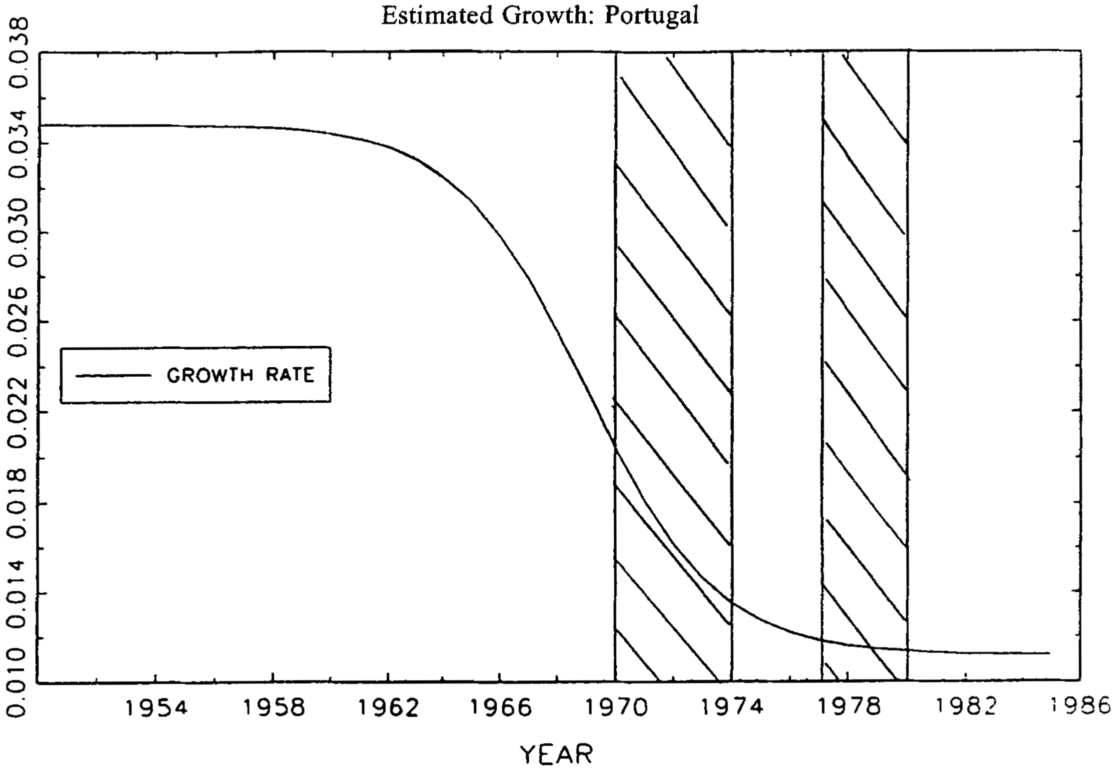


FIGURE 8
Estimated Growth: Spain

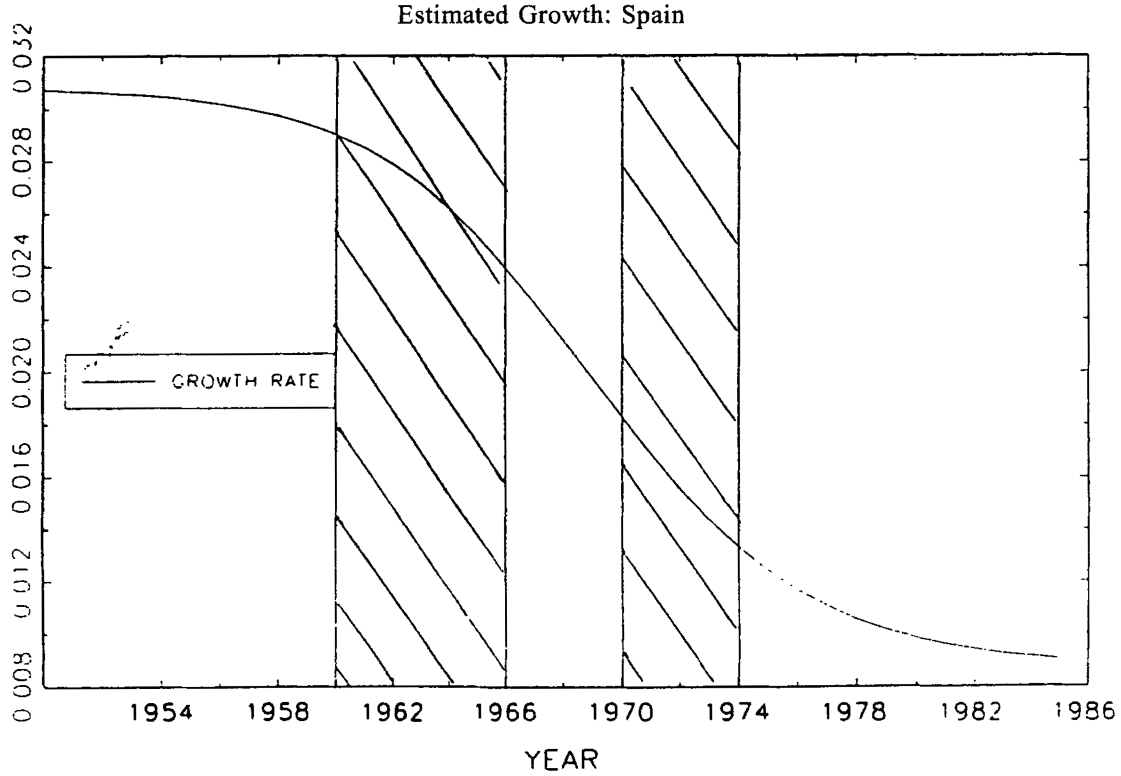


FIGURE 9
Estimated Growth: Sri Lanka

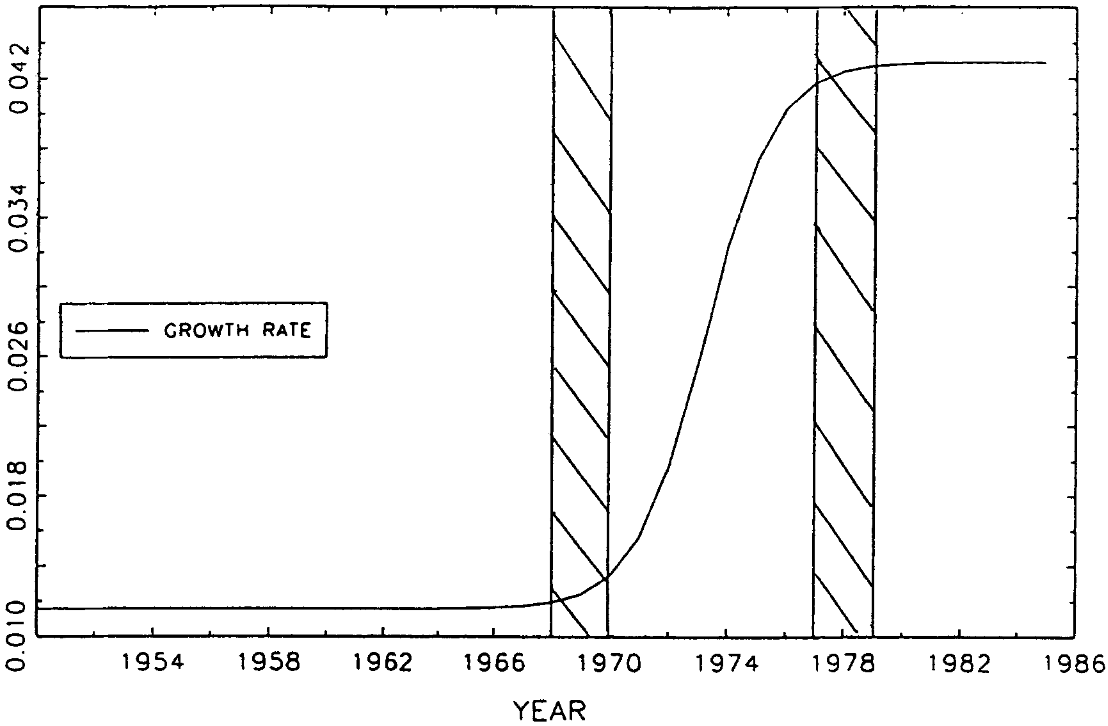


FIGURE 10
Estimated Growth: Argentina

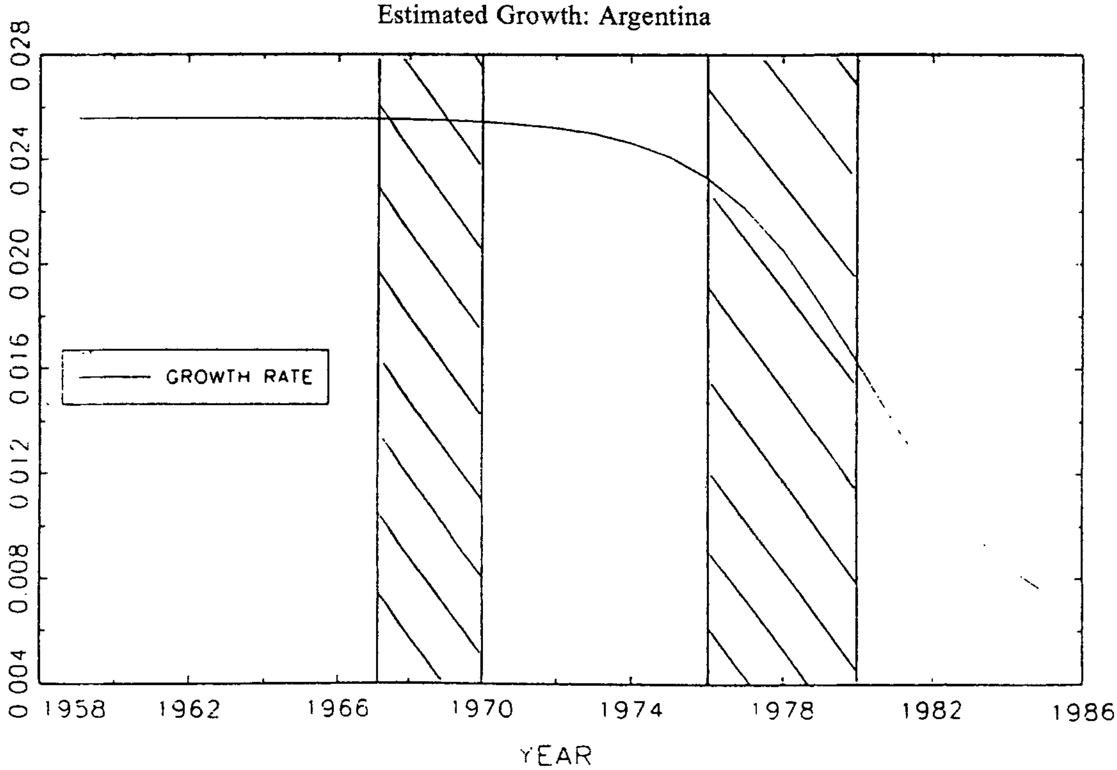


FIGURE 11
Estimated Growth: Yugoslavia

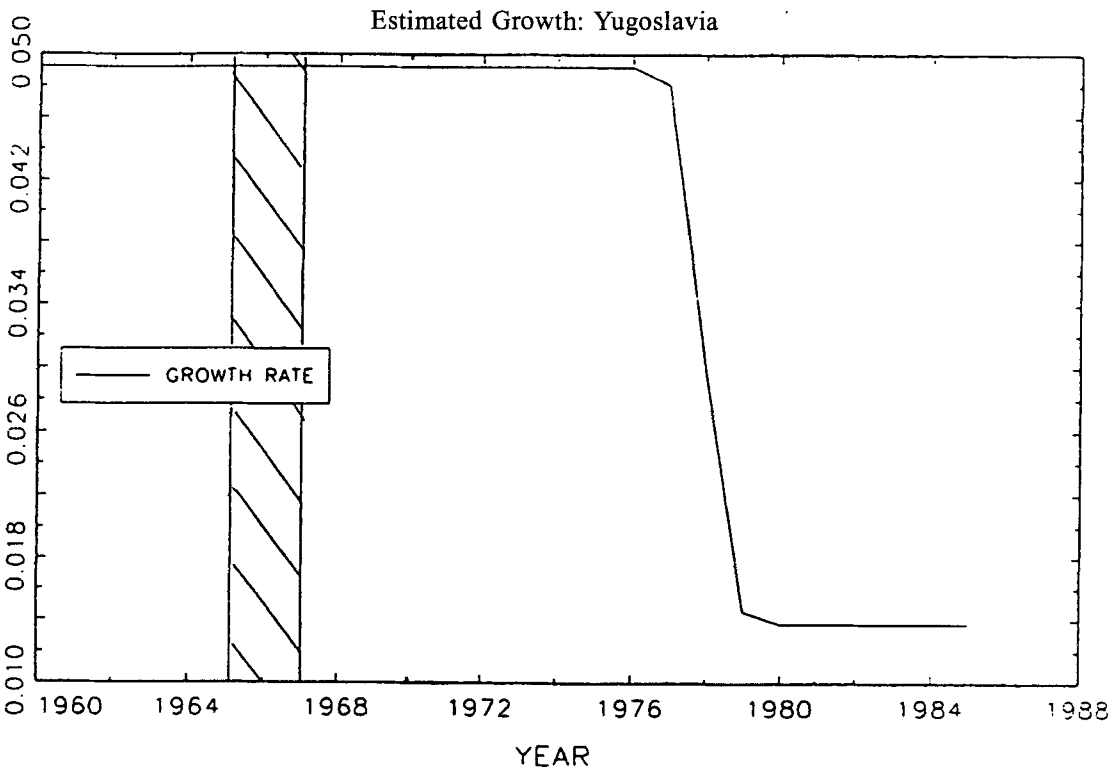
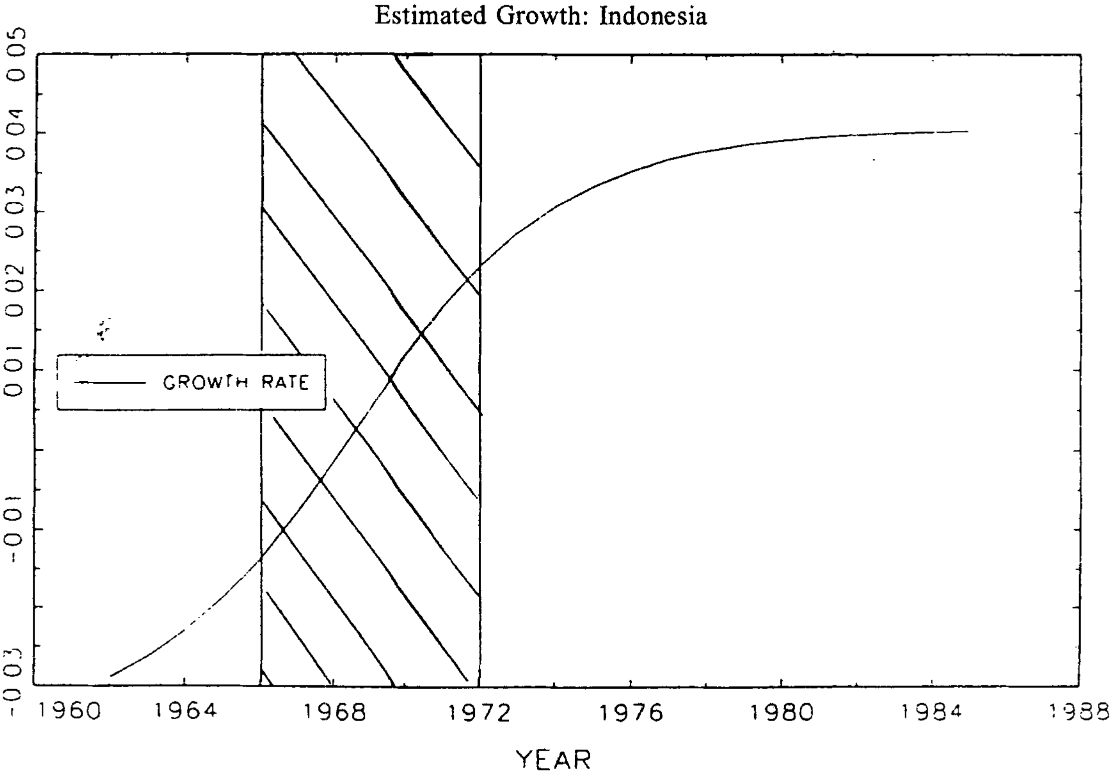


FIGURE 12
Estimated Growth: Indonesia



constitute a formal test of the relationship between liberalization and growth, rather it is an attempt to see whether there are any obvious associations. The range of experience is interesting.

For Columbia (Figure 1) it appears quite possible that liberalization resulted in a stimulus to growth. The start of the trade reforms in the mid-1960s coincides with a clear acceleration in growth of real per capita GDP. For Sri Lanka (Figure 9), the 1968–1970 reforms also appear to have precipitated a rapid acceleration in growth. On the other hand, for Korea (Figure 3) and Indonesia (Figure 12) the fitted models suggest much longer and slower transitions between growth paths have taken place. Indeed, so gradual are the transitions that their effective starting points are estimated to have occurred before the earliest point in the data. The clear implication here is that changes in growth are likely to have had very little to do with liberalization.

In the cases of Greece (Figure 2), Israel (Figure 4), Pakistan (Figure 6) and Argentina (Figure 10) there is evidence that liberalization may have been associated with a fairly sharp decline in the rate of growth towards the end of the episodes. A similar experience is found for New Zealand (Figure 5), although the transition between growth paths here appears to have been fully completed before the end of the 1962–1981 liberalization. For Portugal (Figure 7), Spain (Figure 8) and Yugoslavia (Figure 11) there is somewhat less compelling evidence that the timing of negative transitions in growth rates is related to that of liberalization episodes.

What then, if anything, do our findings suggest? First of all, the very strong connections between liberalization and changes in growth which PMC point to appear to be unwarranted once we model changes in growth as smooth transitions in which the timing of the transitions is data determined, and we compare the fitted transitions to actual liberalizations as identified by the policy accounts of PMC. Second, there is very clearly a rich diversity of experience. This is not surprising, of course, since we would expect liberalization to affect different countries in quite different ways, depending upon inter alia the institutional context within which any reforms are initiated and the degree of slippage from initial commitments which takes place during the liberalization episode. Policy accounts have shown this to be so considerable in some countries that the "liberalization" is essentially a fiction. Third, even where reforms are implemented and slippage is minimal, different economies may be subject to different external shocks during the episode itself.

One can add a little richness to the analysis by probing further the nature of the liberalization episodes themselves. PMC do this by classifying liberalization as strong or weak, depending upon the depth and extent of the trade reforms. They also cross-classify episodes as sustained, partially sustained or collapsed, depending upon the length of time the reforms remain in place. The indicators for the countries for which we have estimates are reported in Table II. Again one sees a rich diversity of experience. The key point to note from the standpoint of the present paper is the lack of any close correlation between strength and duration of liberalization on the one hand and direction and speed of any transition in growth on the other.

Can one say anything about those cases where liberalization *could* be associated with a negative change in growth or is this situation completely implausible? In fact a negative relationship is plausible. Many developing countries "liberalize" under pressure from the multilateral institutions. Although most trade reform conditions are components of World Bank conditionality, the correlation between countries simultaneously undergoing structural adjustment and IMF stabilization programs is high. The latter are typically agreed and implemented in difficult economic circumstances and involve disabsorption. Initial stabilization is often accompanied by a sharp fall in growth, which may appear to be attributable to trade reforms. Again, however, as with possible positive associations one needs to be cautious in ascribing causality, especially since in some cases negative transitions coincide with the global slowdown in growth which occurred in the mid-1970s.

V. CONCLUDING COMMENTS

In this paper we have used a new technique to re-appraise the time-series properties of long-run growth rates of a number of developing countries. Our evidence suggests that (with the exception of Brazil) their growth rate experiences are best described as having

TABLE II
Characteristics of Liberalization in Sample Countries

Country	Episode	Strong	Weak	Sustained	Partially Sustained	Collapsed
Colombia	1964–1966		*			*
	1968–1982	*			*	
Greece	1953–1955	*		*		
	1962–1982		*	*		
Korea	1965–1967		*	*		
	1978–1979		*	*		
Israel	1952–1954	*		*		
	1962–1968	*		*		
	1969–1977	*		*		
New Zealand	1952–1956		*	*		
	1962–1981		*	*		
Pakistan	1960–1965		*		*	
	1972–1978		*		*	
Portugal	1970–1974		*			*
	1977–1980		*		*	
Spain	1960–1966	*			*	
	1970–1974		*		*	
Sri Lanka	1968–1970		*			*
	1977–1979	*		*		
Argentina	1967–1970		*			*
	1976–1980	*				*
Yugoslavia	1965–1967	*				*
Indonesia	1966–1972	*		*		

Source: PMC [1991] vol. 7.

undergone a smooth nonlinear transition through time between two distinct values rather than having stayed constant (as implied by the DFR results), or having switched values instantaneously at any particular point in time.

This is an important finding which challenges the conventional wisdom on changes in growth rates. It also has important implications for the appropriate structural modeling of economic growth: if the data series are best characterized by the kind of smooth transition processes considered here, it follows that any posited structural model must be capable of taking this into account, which means a careful assessment of the statistical properties of any proposed explanatory variables becomes necessary.

To motivate the application of this technique we matched our estimated transitions to data on liberalizations for a (nonrandom) sample of developing countries. There are good theoretical reasons for expecting a causal link between liberalization and growth. With the exception of the PMC study, the empirical evidence is less conclusive. PMC is, however, an important exception in that it has turned out to be very influential. The configuration of our results places a serious question mark against the PMC findings. Not only is there an absence of a complete one-to-one mapping between positive transitions and trade reforms, in some cases the transitions are in fact negative. As we emphasized above, our results cannot be taken as conclusive since we have not conducted a formal test of the impact of liberalization on economic growth. They are, however, strongly suggestive and highlight the important fact that the timing and speed of changes in growth rates need to be treated endogenously.

Finally, the paper points up some interesting directions for future work. First, it is clear that the smooth transition framework could be further developed to capture asymmetries and perhaps also multiple transitions, though both extensions would be rather more demanding in terms of degrees of freedom requirements and estimation effort. Second, structural modeling directed at explaining changes in growth rates resulting from smooth transitions, rather than structural breaks, would be worthy of investigation. Within such a framework the role of liberalization could then be tested in a more formal way.

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