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Financial Development and Economic Growth: An Econometric Analysis for Singapore

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Abstract

With the emergence of the rapidly expanding literature on endogenous growth, the relationship between financial development and economic growth has received a new source of inspiration. Recent cointegration techniques that focus on the estimation and the identification of long-run economic relationship(s) between data variables are particularly appropriate to the study of long run endogenous growth models. This paper has applied these techniques to the Singapore data using a supply-side framework. By and large, the econometric analysis in this paper has yielded results that are in line with predictions of endogenous growth models. In particular, we find that financial development positively affects both transitional and long-run growth in Singapore.

JEL Classification: C50, 042

1/ This paper is a revised version of chapter four of my Ph. D. thesis. I wish to thank my dissertation advisers David Hendry, John Muellbauer, and Christopher Adam, all at Oxford University, for their insightful commentary on earlier versions of the paper. I am also grateful to seminar participants at Nuffield College, Oxford and colleagues at the Fund for comments. The usual disclaimer applies.

Contents

Page

Charts

Summary

The relationship between financial development and economic growth has remained a topical issue in development economics. New endogenous growth models have studied financial factors in an attempt to analyze formally the interactions between financial markets and economic growth. Recent modeling of the impact of credit on the aggregate economy has emphasized the endogeneity of the financial structure itself. These models study how the level of financial development is itself endogenously linked to the aggregate level of activity. Drawing on the theoretical foundations of endogenous growth and endogenous financial structure models, this paper employs the vector autoregression (VAR) econometric techniques to analyze the impact of financial development on economic growth in Singapore. The econometric models developed in the paper use quarterly time series data over the period 1975Q1 to 1991Q2.

An aggregate production function for Singapore is specified that consists of the basic factor inputs of labor, physical capital, and the stock of human capital and is augmented by a financial development variable. Vector autoregression cointegration techniques are used to estimate and identify long-run equilibrium relationships between financial development and growth in the augmented production function. The estimated long-run cointegrating solutions are then combined with short-run factors of economic adjustment in a bivariate system to form error-correction (ECM) growth models for Singapore.

The econometric results of the estimated augmented aggregate production function for Singapore in this paper are generally in line with predictions of endogenous growth theory. In particular, they indicate that growth in Singapore is driven mainly by endogenously determined variables, among which is financial development. Moreover, financial development is found to influence long-run growth in Singapore largely by its positive effect on investment efficiency.

I. Introduction

Financial liberalization and the development of the financial sector have been assigned strategic importance in economic development. In many countries of Asia, Latin America, and much more recently, the transition economies of Eastern Europe, financial sector reforms remain high on the agenda of policymakers. Structural adjustment programs undertaken by many developing countries in collaboration with the IMF and The World Bank now typically include financial sector reforms designed to increase the extent to which savings become available and the intermediation of these savings to investment opportunities that bring the highest return.

During the past decade nearly all the countries in the Pacific Basin Region liberalized their financial systems. In Indonesia, Malaysia, and the Philippines, significant reforms began between the early to mid-1980s. More gradual steps toward liberalization, which were intensified in the latter half of the 1980s, occurred in Korea and Thailand. Nepal began to implement a financial liberalization program in the second half of the 1980s.

In contrast, Singapore had largely liberalized its financial system by the mid-1970s. The potential for the financial sector to become a growth sector, serving the needs of not only the domestic economy but also the regional and international economy, was recognized in the late 1960s. The Monetary Authority of Singapore (MAS) explains the financial development strategy mapped in the late 1960s as follows:

"The financial sector was to be developed as a major growth industry in its own right, rather than fulfill a subsidiary role to meet the needs of the other sectors of the economy. The policy was to expand the role of the financial sector beyond its traditional functions. The aim was to develop it into a modern sophisticated financial center to serve the needs not only of Singapore and the surrounding region but also beyond. The main potential benefits of this policy were perceived to be the increase in the flow of trade and investment and the economic growth of Singapore and the region..." (Monetary Authority of Singapore (MAS) Annual Report (1981), p. viii).

With these objectives in mind, cautious financial policies were undertaken to improve investor confidence from both inside and outside, which has not only led to an increased monetization but also greater capital deepening of the economy. These efforts contributed to the rapid growth of the financial sector, which became the second largest contributor to both GDP (gross domestic product) and employment in the economy after manufacturing. Contributing 16.6 percent of GDP in 1970, in 1991 the Singapore financial and business sector accounted for 25.5 percent of GDP, a sizable proportion even by the standards of developed countries.

Consequently, by the early 1980s, when other countries in South-East Asia were beginning to institute financial sector reforms, Singapore already had a relatively well-developed financial sector that was closely integrated with international financial markets. This paper examines the role of financial development in the economic growth of Singapore, using an aggregate production function approach. 1/

From a more technical perspective, the availability of macroeconomic data allows us to use more recent techniques in the analysis of cointegrating time series relationships developed by *inter alia* Johansen and Juselius (1990) and do a more comprehensive analysis of the growth effects of financial development in Singapore. The results of these methods applied elsewhere by Adam (1992) on Kenya, Ahumada (1992) on Argentina, Baba, Hendry and Starr (1992) on the United States, Hendry and Mizon (1990, 1993) on the United Kingdom and Johansen and Juselius (1990) on Denmark and Finland and others, suggest that they can be profitably applied to a South-East Asian country. Archetypical evidence from the Singapore economy, now considered as a developed nation, will not only provide us with the opportunity to develop a rich understanding of factors that have governed the growth experience of a Newly Industrializing Economy (NIE), but will also provide some indication of the likely evolution of macroeconomic aggregates in other countries in South-East Asia where the process of financial liberalization and financial development started much later.

The analysis begins in Section II with a specification of an aggregate production function, which consists of the basic factor inputs of labor, physical capital, and the stock of human capital, and is augmented by a financial development variable. Endogenous growth models emphasize factors that determine the long-run growth path of an economy and thus cointegration techniques can be profitably applied to them. Section III studies the impact of financial development on the long-run growth path of the Singapore economy by examining the cointegration properties of the estimated augmented aggregate production function. Weak exogeneity is tested in this section and found to be invalid in the augmented aggregate production function, largely because of the cross-equation links between output and labor. Section IV is therefore devoted to estimating a bivariate system model (output, labor) of the augmented aggregate production function. This section employs non-nested encompassing tests in order to discriminate between the competing estimated growth models for the Singapore economy. Section V examines the statistical robustness of the dominant estimated aggregate production function growth model for Singapore and then presents its economic interpretation. Finally, Section VI concludes and suggests some areas where this work could be extended. The appendices contain a formal description of the cointegration methodology employed in the paper and the data description and sources.

1/ Chapter 3 of the author's dissertation examines this issue through an alternative approach, namely in an augmented Keynesian aggregate demand system.

II. Specification and Estimation of the Long-Run Aggregate Production Function

The economic growth of a nation can, in general, be attributed to three factors: growth in capital, growth in labor, and technical progress (including increases in efficiency). In the neoclassical growth models, growth in the long run is tied down by exogenously determined variables, for example technical progress. In contrast, endogenous growth theories seek to discover what lies behind the exogenous rate of technical progress and hence a country's growth rate. The key difference is that these models require constant or increasing returns to scale. This then implies that those variables that lead to the non-decreasing returns drive the growth rate. Numerous candidates have been suggested as the source of non-decreasing returns: in particular, the stock of human capital in various guises, Lucas (1988); accumulated capital, Rebelo (1990); research and development, Romer (1986, 1987, 1990); or public infrastructure investment, Barro (1991). Thus, endogenous growth models highlight sectors of the economy that influence the growth path of an economy.

In their study of the sources of growth in the Newly Industrializing Economies (NICs) on the Pacific-Rim (Singapore included), Kim and Lau [1993) concluded that by far the most important source of economic growth in the NICs is capital accumulation, accounting for more than 80 percent of their growth rate. This contrasts with their results for the G-5 Industrialized countries, in which technical progress played the most important role. In a comparative study of Hong Kong and Singapore, Young (1992) also concluded that Singapore's high growth rate is largely due to the rate of physical capital accumulation. The natural question to ask, then, is to what extent has the financial sector contributed to this capital-oriented growth documented for the NICs. More precisely, are there reasons to believe that the financial sector is one of the determinants of the growth path of an economy?

Financial development has a dual effect on economic growth. On the one hand, the development of domestic financial markets may enhance the efficiency of capital accumulation; on the other hand, financial development can contribute toward raising the savings rate and thus the investment rate, and lead to more capital accumulation. 1/ The latter point was made by McKinnon-Shaw (1973), while the former channel was emphasized by Goldsmith (1969), who also documented positive correlation between financial development and the level of real per capita GNP.

1/ This may not always follow, since they could choose to invest their savings in foreign assets. Moreover, in the U.K context, increased financial intermediation led to the lowering of the savings rate in the 1980s.

More formally, consider a Rebelo (1991) type production function which for simplicity, is assumed to depend only on the capital stock,

$$y_t = Af(k_t) \qquad (1)$$

where y_t and k_t denote output and the physical capital stock respectively, at time t, and A is a constant. By totally differentiating equation (1) and denoting the rate of growth of output by Δy, the savings rate (dk/y) by s, and the marginal productivity of physical capital by ϕ, we have:

$$\Delta y_t = A \frac{dk_t}{y_t} f'(k_t) \equiv As_t\Phi_t \qquad (2)$$

From equation (2), the rate of output growth is the product of the savings rate (s_t) and the marginal productivity of capital (ϕ_t). While the McKinnon-Shaw (1973), Kapur-Mathieson (1981, 1980a) models of the effect of financial development on growth largely work through the savings (s_t) variable, the Goldsmith (1969) effect works through the productivity (ϕ_t) variable. The new endogenous growth models tend to emphasize the latter hypothesis and thus focus on the dominant effect of financial markets on growth through its impact on ϕ_t.

Cointegration techniques focus on the estimation and the identification of long-run relationships between data variables and are therefore particularly appropriate to the study of long-run growth. Using these techniques within a Cobb-Douglas aggregate production function but augmented with a trade variable, Coe and Moghadam (1993) established that trade and capital, broadly defined, account for all the growth in the French economy during the past two decades. We apply the same principles here to examine the extent to which financial development is a determinant of the long-run growth path of the Singapore economy. This is achieved by studying the cointegration properties of an aggregate production function in which financial development (proxied by credit) 1/ is embedded. The argument is that, the growth rate in the long run, is determined by supply-side variables, namely, labor, physical capital, and human capital, together with a financial development variable. Since financial development is proxied by credit, the approach in this paper is thus analogous to embedding money in a production function. More formally the augmented aggregate production function can be specified as,

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1/ Admittedly, there are other measures of financial development such as real interest rates or the level of monetization in an economy or output of the financial sector. Chapter 3 of the author's dissertation discusses the strengths and weaknesses of these four measures of financial development and then evaluates them on the Singapore data using non-nested encompassing tests. The results of this exercise show that the optimal measure of financial development in Singapore is the credit proxy (fdc).

$$y_t = \phi(fdc_t)^{\delta_1} hc_t^{\delta_2} k_t^{\alpha} l_t^{\beta} \qquad (3)$$

where y is output, k is physical capital, l is labor, hc is human capital per head, and fdc is the financial development variable with $\phi'(\mathrm{fdc}) > 0$. Following the discussion above, we shall think of the financial sector as increasing the microeconomic efficiency of the whole macroeconomy; in particular, it contributes to an efficient allocation of the physical capital stock.

In following this approach we have to make a choice of a functional form. In the theoretical literature Cobb-Douglas is the norm. Some work has been done on this for Singapore (Chen (1993) and Kim and Lau (1993)). Using the two factor inputs of physical capital and labor, both Chen (1993) and Kim and Lau (1993) rejected the Cobb-Douglas functional form for the Singapore aggregate production function. However, while paying due attention to these results, it is worth noting that our aggregate production function above is a more general one. It augments the basic factor inputs, physical capital and labor with human capital and financial development variables. Within this more general aggregate production function, we estimate a long-run regression equation of the form,

$$y_t = \alpha + \alpha_k k_t + \alpha_l l_t + \delta_1 fdc_t + \delta_2 hc_t + \epsilon_t, \qquad (4)$$

where all the variables are expressed in logarithmic terms and ϵ_t is the random term. 1/ The graphs of the various variables are shown in the next few pages and all the data series employed in this paper are seasonally unadjusted. The reasons for this are twofold. First, in modeling certain variables that are obviously seasonal in nature, such as growth, we may be actually interested in the seasonal pattern in the data for forecasting purposes. Second, as pointed out by Wallis (1974), using seasonally adjusted data can distort the dynamics of the estimated model, alter the exogeneity status and hence lead to wrong inferences. This is reinforced by the fact that seasonal adjustment procedures often create a bias toward a non-rejection of the unit-roots hypothesis (Perron and Vogelsang (1992)). It is thus preferable to use seasonally unadjusted data in applied work.

1/ Our capital stock data does not include R & D capital expenditure, even though R & D is an important variable in the new endogenous growth models. This is partly due to lack of data but more significantly it is because R & D is not usually capitalized but expensed. See also the Canada Consulting Group (1992) pp 6-14.

Chart 1. Real Income (y), Physical Capital (k), and Financial development (fdc) (in logarithmic terms).

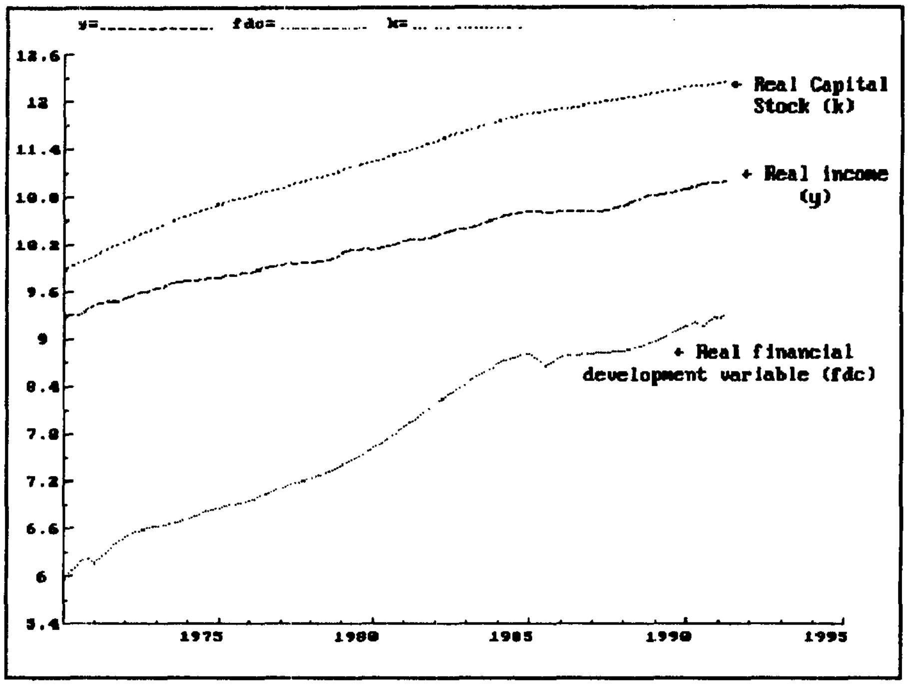


Chart 2. Labor- l (employment per total hours worked)

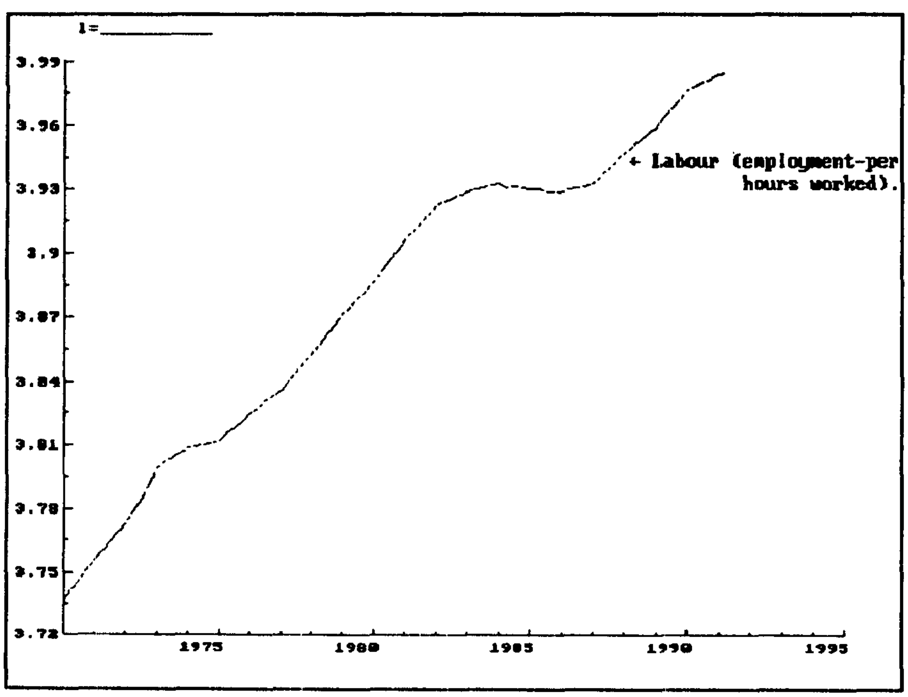
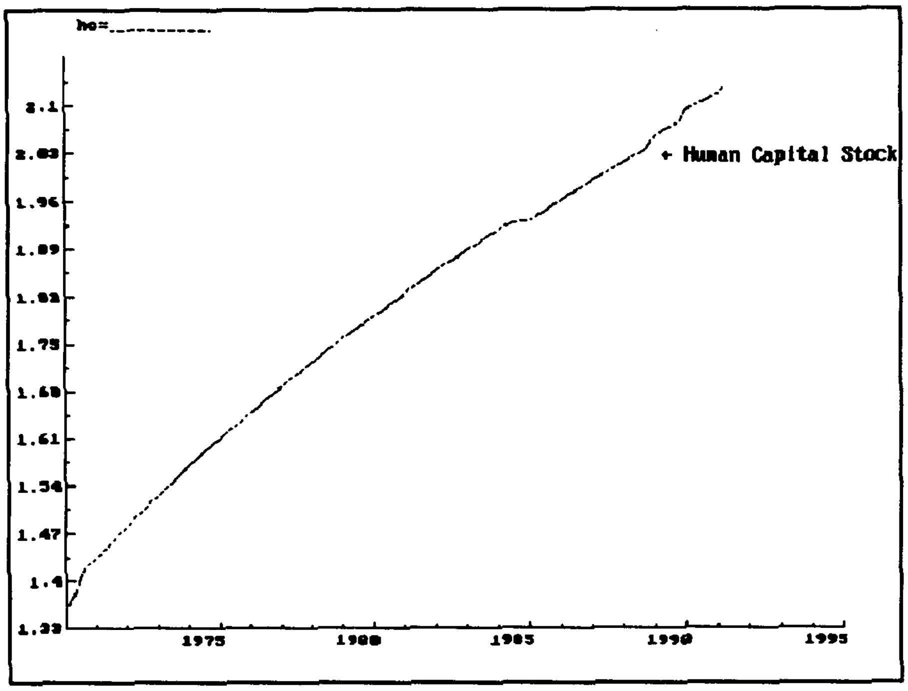


Chart 3. Human Capital Per Head (hc).



Before we test for a cointegrated long-run relationship between the variables of the augmented aggregate production function, we first establish the order of integration of the individual variables. The table on the next page reports the unit-roots test results for the variables using the Augmented Dickey-Fuller (1981) test (ADF), the Sargan-Bhargava (1983) cointegrating regression Durbin-Watson statistic (CRDW), and the Phillips-Perron (1988) Z-statistics. The low power of all unit-roots test necessitates this and consequently the decision on the order of integration of a variable will be based on the weight of evidence emerging from the three tests.

The critical value for a sample size N = 100 when the ADF regression includes a constant and a trend is -3.45 and -2.89 when only a constant is included, both at the 5 percent significance level. The corresponding critical values at the 1 percent level of significance are -4.04 and -3.51 respectively (Banerjee et al, (1993)). For all variables except real income (y), the latter is used as the relevant critical value, as the trend is significant only in the ADF regression for real income (y). The CRDW statistics are compared with the critical values in Engle and Yoo (1987) and this is equal to 0.39 for a sample size N = 100 and at the 5 percent significance level and 0.51 at the 1 percent significance level. The Z statistics are compared with the critical values in Phillips and Ouliaris (1987). In computing the Z-statistics, four auto-covariances were used as in Phillips and Perron (1988). The ADF regressions and the CRDW values were estimated using PC-GIVE (7.00) and the Z-statistics were computed in GAUSS 386 (COINT). In the tables below, ** and * denote significance at the 1 percent and 5 percent levels, respectively.

Table 1. Unit-Roots Test Results of the Individual Data Variables (ADF and CRDW Tests) -Sample Period 1970Q1-1991Q2

VARIABLE	ADF coefficient	T-STATISTIC	LONGEST LAG	CRDW
y	-0.1876	-2.078	4	0.010
Δy	-0.6213	-3.296	3	1.972**
k	-0.0345	-2.107	5	0.047
Δk	-0.2087	-3.831**	4	2.321**
l	-0.1143	-1.921	6	0.227
Δl	-0.3007	-3.965**	3	2.219**
fdc	-0.1116	-2.621	4	0.116
Δfdc	-0.2431	-3.546*	3	1.998**
hc	-0.4212	-3.323*	5	0.372
Δhc	-0.0921	-4.591**	4	2.110**

Table 2. Phillips and Perron (1988) $Z(\underline{\alpha})$ and $Z(t_{\underline{\alpha}})$ Statistics- Sample Period 1970Q1 - 1991Q1

VARIABLE	$Z(\alpha)$	$Z(t_\alpha)$
y	-1.992	-0.687
Δy	-21.10**	-13.38*
k	-2.045	-1.339
Δk	-26.32**	-17.49*
l	-1.732	-0.764
Δl	-30.80**	-19.56*
fdc	-2.226	-1.247
Δfdc	-29.71**	-18.02*
hc	-1.839	-0.665
Δhc	-28.30**	-16.61*

With the exception of the human capital variable, which turns out to be a stationary I(0) series from the Augmented Dickey-Fuller (ADF) test, all the data series appear to be integrated of order one, that is I(1). As the recursively computed ADF t-statistics for the one-period lagged dependent variable in the ADF regression for the various variables indicate in the next pages (Charts 4 - 8) 1/, physical capital (k), labor (l), and financial development (fdc) are clearly non-stationary I(1) series, while the human capital variable is a stationary I(0) series. However, both the cointegrating Durbin-Watson statistics (CRDW) and the Phillips-Perron (1988) Z-statistics identify the human capital variable, as an I(1) series. 2/ From the ADF test, the income term y is a borderline I(1), I(2) variable, but the recursively computed ADF t-statistics for Δy_{t-1} variable, shown in Chart 4 above reveal these be highly variable so that both reject and non-reject decisions occurred. In fact, Δy_t tends to be a stationary I(0) series for a large part of the sample period, which implies that y_t is, on the whole an I(1) series. In general, inference from unit-roots tests in small samples is not always reliable. In particular, structural breaks and regime shifts are likely to bias such tests in favor of a unit-root. Therefore, tentatively we accept that all the five variables including human capital are I(1) series, while expecting the cointegration analysis to reveal further stationarity features of the data variables in the VAR space.

1/ The recursively computed ADF t-statistics allow us to assess whether the null hypothesis of a unit-root can be maintained over the sample as a whole.

2/ The fact that the deterministic trend is not significant in the hc_t ADF regression could potentially be the explanation for this counter-intuitive unit-roots test result for human capital. The use of deterministic trends when the trend is actually stochastic can bias these test against the null hypothesis of a unit-root (Harvey et al (1986)). However, in the final system equation estimation of the augmented production function, our human capital variable is represented by Δhc_t rather than hc_t in order to circumvent any potential problem that may emerge owing to this counter-intuitive result from the ADF unit-roots test.

Chart 4. The Recursively Computed ADF t-statistics For the Variable Δy_{t-1} in the ADF Regression.

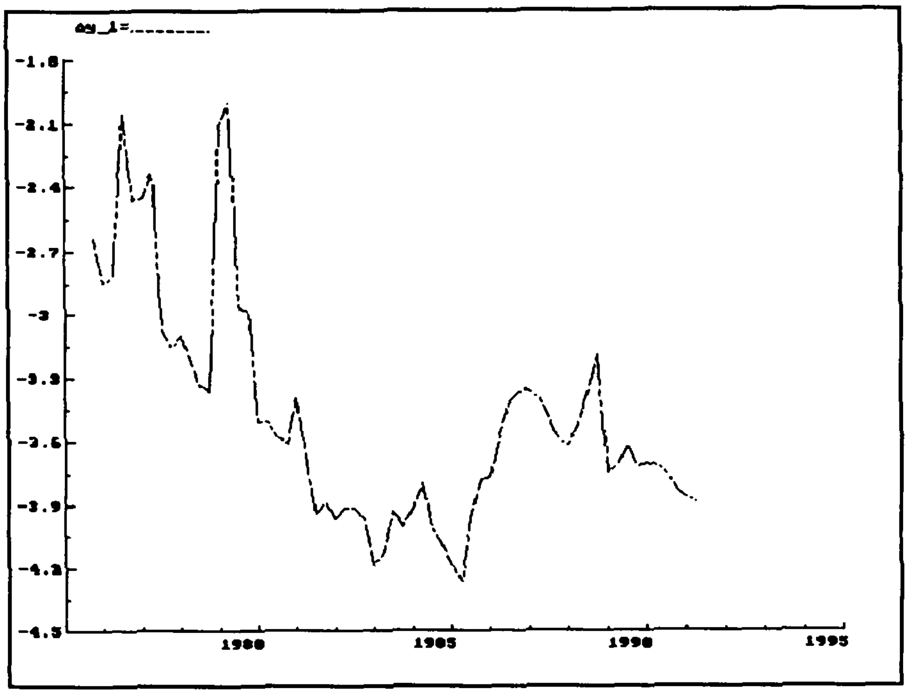


Chart 5. The Recursively Computed ADF t-statistics For the Variable k_{t-1} in the ADF Regression.

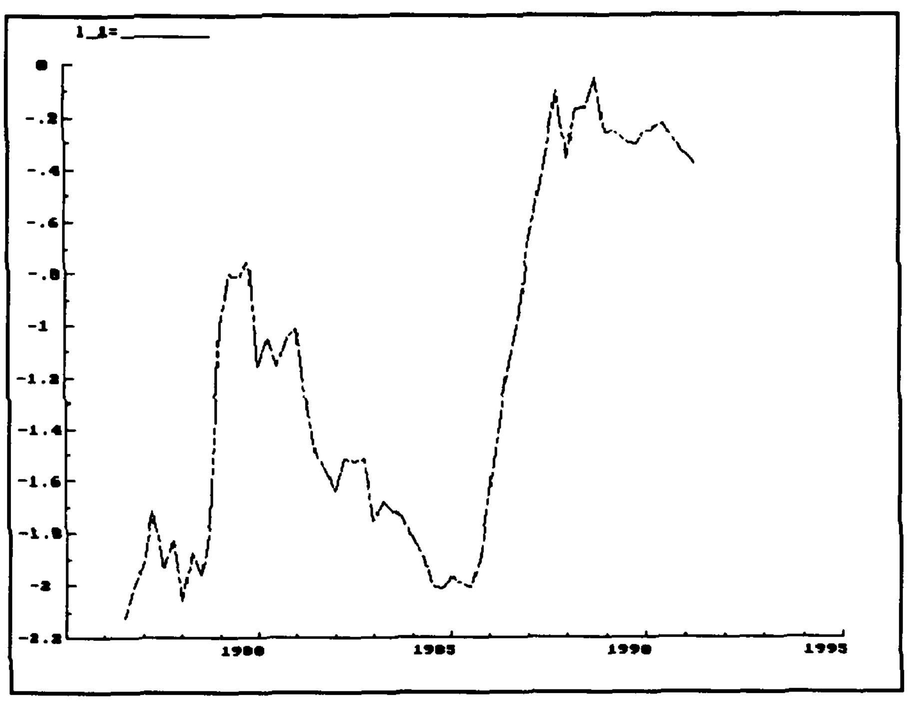


Chart 6. The Recursively Computed ADF t-statistics
For the Variable l_{t-1} in the ADF Regression.

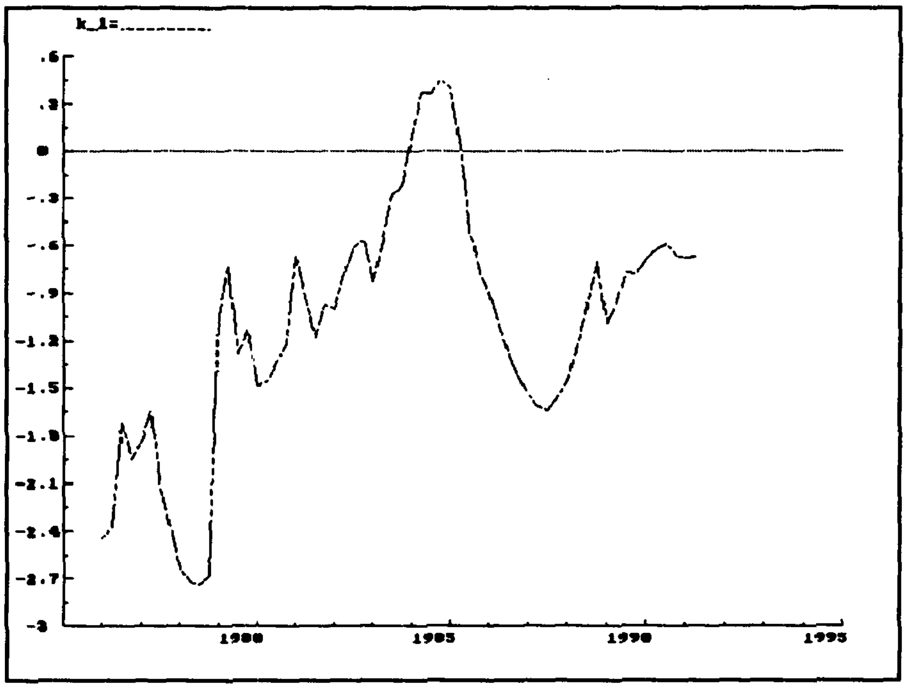


Chart 7. The Recursively Computed ADF t-statistics
For the Variable fdc_{t-1} in the ADF Regression.

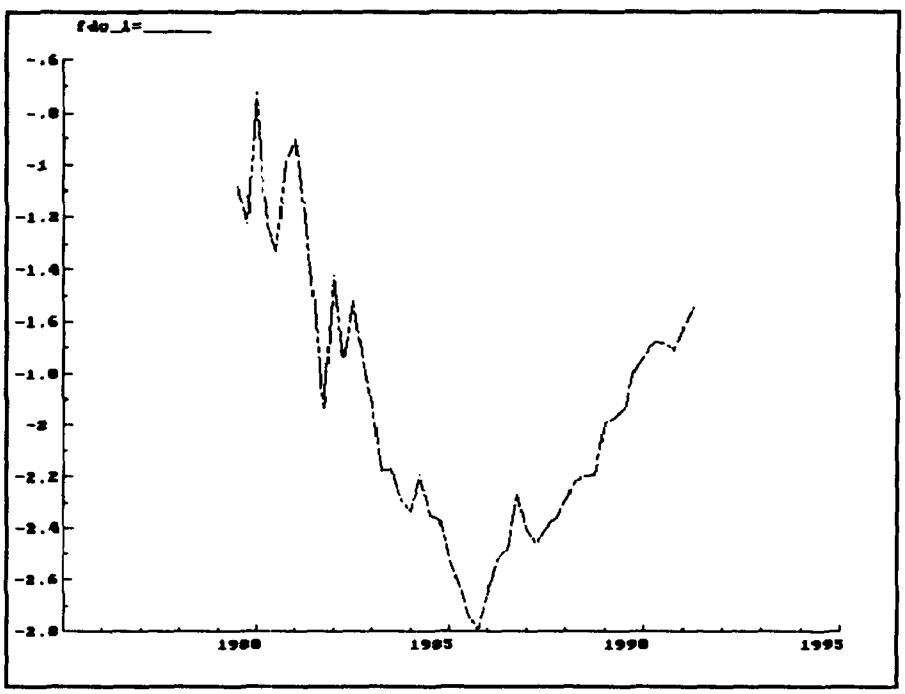
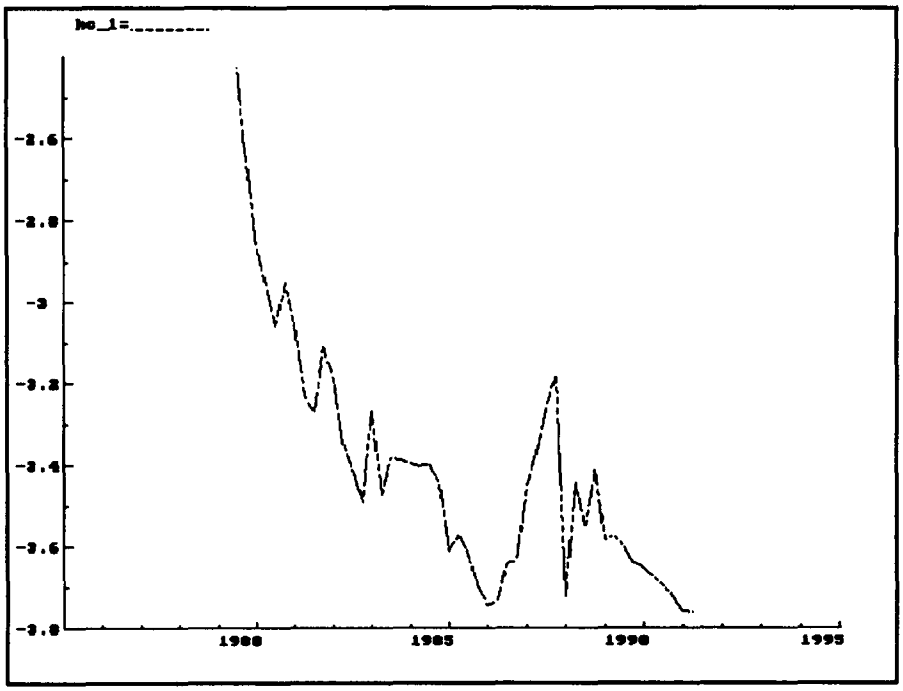


Chart 8. The Recursively Computed ADF t-statistics
For the Variable hc_{t-1} in the ADF Regression.



Having established the order of integration of the individual data series 1/, we now proceed to test for cointegrating relationship(s) among them if any. Following this approach, we commence by estimating a vector autoregression (VAR) model with a constant and a trend. 2/ The VAR is estimated with four lags on each variable; all the diagnostic tests identified this to be the optimal lag length. The three evaluation statistics are tests for vector autocorrelation of 5^{th} order (F^{v}_{ar} (125, 127)), a system heteroscedasticity test (F^{v}_{het} (780, 34)), which here exhausts the degrees of freedom and has no power, and a test for multivariate normality ($\chi^{2v}_{nd}(10)$ (Hendry and Doornik (1993)). The 1-step Chow statistics graphs indicate parameter constancy for all the equations of the VAR based on this lag length. The baseline innovation variances are small and reveal no obvious problem. The dynamics of the system were investigated by orthogonalized impulse response analysis; although the impulse response graphics are not reported here, the contemporaneous responses indicate that the contemporaneous correlations of the variables are insignificant. Moreover, most of the variables in the VAR show the expected theoretical prior reactions to a positive unit innovation shock in the equation standard error. Thus a constant parameter congruent Unrestricted VAR has been obtained (Clements and Mizon (1991)). The estimated results of this congruent Unrestricted VAR (UVAR), using the Johansen multivariate cointegration analysis (Johansen and Juseliues (1990)), are shown in Table 3 on the next page. The sample period extends from 1971Q1 to 1991Q2. Though containing 82 observations, this sample is relatively short and, as is characteristic of high frequency data estimated over short periods, it may be subject to a low signal-noise ratio, thus requiring careful interpretation of the statistics. 3/

1/ All the variables of the aggregate production function were tested for the presence of seasonal unit-roots in them, using the Augmented Dickey-Fuller (1981) version of the seasonal integration test (Charemza and Deadman (1992)). The results of this test indicate that the production function variables are seasonally stationary.

2/ The VAR is estimated using version 7.01 of the PC-FIML software by Hendry and Doornik (1994).

3/ The relatively high values of the eigenvalues μ_i, as reported in Table 3, however reveal that the signal-noise ratio is not that low. On the role of the signal-noise ratio in cointegrated systems, see Kostial (1993).

Table 3. The Johansen-Juselius (1990) Procedure: VAR with Four Lags, a Constant and a Trend- Sample Period: 1971Q1 - 1991Q2

Diagnostics: **Goodness of fit and evaluation statistics**

	y	k	l	fdc	hc
σ	0.013	0.012	0.034	0.020	0.003
$r(y,\hat{y})$	0.997	0.996	0.998	0.998	0.999

$F^{v}ar$	(125,127)	1.23
$F^{v}het$	(780,34)	0.39
$\chi^2 vnd$	(10)	8.1

The test statistics: 1/ 2/
(Testing the number of cointegrating vectors)

Test	ρ=0	ρ≤1	ρ≤2	ρ≤3	ρ≤4
μ_i	0.478	0.377	0.267	0.150	0.099
Trace	139.5**	86.15**	47.35	21.88	8.55
5%CV	87.3	63.0	48.4	25.3	12.2
μmax	53.31**	38.80*	25.47	13.33	8.55
5%CV	37.5	31.5	25.5	19.0	12.2

1/ Trace is equivalent to $-T\sum\log(1 - \mu_i)$.
2/ μmax is equivalent to $-T\log(1 - \mu_i)$.

The eigenmatrix, β'

β'	y	k	l	fdc	hc	$Trend$
1	1.00	-0.778	-0.312	-0.412	-0.087	0.005
2	-3.682	1.00	1.74	0.95	0.415	0.009
3	-6.40	-5.725	1.00	-1.029	-0.018	0.024
4	-2.784	0.014	0.51	1.00	3.432	0.003
5	-0.562	0.054	0.025	0.054	1.00	0.002

The adjustment coefficients α corresponding to the cointegrating space

y	-0.095	0.1967
k	-0.044	-0.036
l	0.058	-0.254
fdc	-0.037	0.039
hc	-0.023	0.0026

In the matrix showing the test for the number of cointegrating vectors, test statistics marked with * and ** mean that the null hypothesis of no cointegrating vectors is rejected at the 5 percent and 1 percent levels, respectively. From this matrix of test statistics and the critical values tabulated by Osterwald-Lenum (1992), there are two cointegrating vectors according to both the "trace statistics" and the "maximal eigenvalue" test, μmax. The first and second significant cointegrating vectors (standardized on the principal diagonal) are represented in row 1 and row 2 of the β' eigenmatrix, respectively. Moreover, here the trend is found to be significant in the production function system and is therefore restricted to lie in the cointegration space, as shown in the β' eigenmatrix. Across both cointegrating vectors, there is no evidence of any of the variables being stationary I(0) series, which therefore reinforces the broad conclusion reached by our unit-roots tests.

III. Evaluating the Long-Run Cointegrating Vectors: Data Robustness and Economic Interpretation

In the first cointegrating vector, all the estimated coefficients have the expected theoretical signs and are of reasonable magnitudes, although the normalized coefficients for physical capital and labor are higher and lower respectively, than their factor shares. 1/ The elasticity of output with respect to physical capital in this cointegrating vector is larger than that reported by Chen (1993). Across the first cointegrating vector, the estimated coefficients on capital (k), labor (l), the financial development variable (fdc), and human capital (hc) sum to 1.589, thus implying increasing returns to scale (IRS) with respect to the factors in the long run. This conforms to the prediction of Rebelo (1990) type *AK* models of endogenous growth and is in accordance with the result obtained by Kim and Lau (1993) in their estimated aggregate meta-production function for four Newly Industrializing Economies (NIEs). 2/ Moreover, in Romer's (1986) work on long-run growth, this condition is necessary for endogenous growth. This is because a sum of the factor coefficients of less than 1 implies decreasing returns which, in the absence of exogenous technological progress, implies zero steady state growth. The *AK* models of endogenous growth emphasize that a marginal product of accumulable factors bounded substantially away from zero (see Jones and Manuelli (1990)) would predict that any economy experiencing rapid factor income accumulation should show a rapid decline in the share of the national income accruing to the principal nonaccumulable factor - labor. However, almost all factor share calculations done for Singapore (see, for instance, Tsao's work in Chong-Yah and Lloyd (1986) p. 21), and Young (1992) show a remarkable constancy of the share of unskilled labor in total factor payments.

Short-run diminishing returns caused by adjustment costs are consistent with long-run constant returns to scale (CRS) with respect to capital assumed in the *AK* models of endogenous growth (see Young (1992)). In the first cointegrating vector, this implies long-run homogeneity of output with respect to capital. In order to test the validity of this hypothesis across the first cointegrating vector, we can impose the restriction of the form,

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1/ The estimated sample average of factor shares for physical capital and labor in Singapore are 0.6 and 0.4, respectively (Chong-Yah and Lloyd (1986) p. 21). However, endogenous growth theories have pointed out that the coefficient on the physical capital may be much larger than suggested by its factor share in an imperfectly competitive environment (Romer (1990) and Rebelo (1990, 1991)).

2/ In fact, the restrictions that the estimated coefficients on physical capital and labor are equal to the factor shares, and that the sum of the coefficients on physical capital, labor, human capital, and financial development sum to unity, are both rejected across the first cointegrating vector.

$$H_0: \beta = C\phi,$$

where C is a p × s matrix of constraints and ϕ the corresponding s × v matrix of unrestricted parameters. C is a 6 × 5 matrix and of the form,

$$\begin{pmatrix} 1.0 & 0 & 0 & 0 & 0 \\ -1.0 & 0 & 0 & 0 & 0 \\ -0.3 & 1 & 0 & 0 & 0 \\ -0.4 & 0 & 1 & 0 & 0 \\ -0.1 & 0 & 0 & 1 & 0 \\ 0.0 & 0 & 0 & 0 & 1 \end{pmatrix} \tag{5}$$

and ϕ is a 5 × 5 matrix. The likelihood ratio test of the significance of this restriction is given by,

$$-2\ln(Q:H) = T\sum \ln\left(\frac{(1 - \mu*_i)}{(1 - \mu_i)}\right) \tag{6}$$

which is distributed as χ^2 with v(p-s) degrees of freedom (Johansen and Juselius (1990)). For the maximal eigenvalues this gives -2ln(Q:H) = 17.64 compared with a critical value of χ^2 v(p-s) = $\chi^2_{0.05}(5)$ = 11.07, thereby rejecting the restriction.

The coefficient on the trend in the first cointegrating vector is negative, implying negative long-run total factor productivity. If we interpret the trend as capturing technical progress, this means that technical regress has occurred in Singapore during this sample period. 1/ Lack of technical progress in a growing economy is consistent with the increasing returns to scale (IRS) result in the first cointegrating vector.

The second long-run cointegrating vector, normalized on the capital stock k in the second row of the eigenmatrix β' is not directly amenable to economic interpretation. However, if we apply a simple row reduction in the cointegrating matrix space and normalize the second cointegrating vector on the output variable (y) we get,

$$\begin{aligned} y &= 0.272k + 0.473l + 0.258fdc \\ &+ 0.113hc + 0.002Trend \end{aligned} \tag{7}$$

1/ Indeed, Kim and Lau (1993) found a negative capital-augmenting technical progress for both Singapore and South Korea.

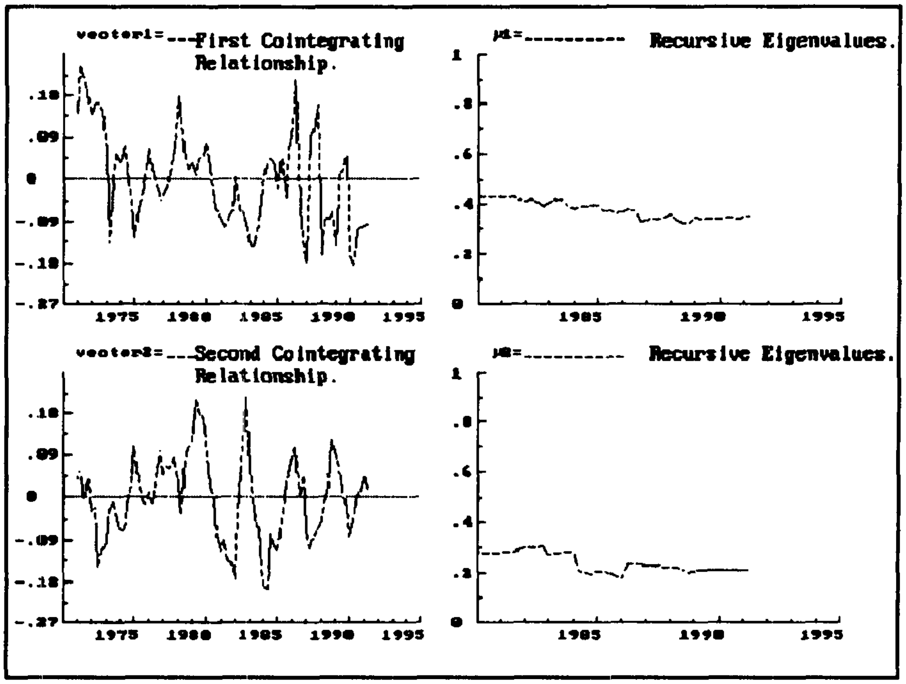
which has features of a long-run output equation. It has the right signs for the factors, and the estimated coefficients for physical capital, labor, financial development, and human capital sum to 1.116, thus satisfying the Cobb-Douglas specification in the limit. The concentrated cointegrating graphics corresponding to these two significant cointegrating relationships and their respective recursively computed eigenvalues (μ_{is}) are shown on the next page.

In contrast to the first cointegrating vector, the output elasticity with respect to physical capital in the second cointegrating relationship is well below its factor share. However, in an imperfectly competitive environment, the interpretation of the factor elasticities with respect to output needs some caution. Since both implicit and explicit subsidies play an important role in Singapore, subsidies to physical capital accumulation can cause the true elasticity of output with respect to physical capital to be less than its observed factor share. By the same line of reasoning, since labor in Singapore is taxed through both the employee's contribution to the Central Provident Fund (CPF), a national pensions scheme, and the Skills Development Fund (SDF), the coefficient on labor may be slightly higher than its factor share, as evident in the second cointegrating vector. More significantly, the addition of the elasticity value for physical capital (k) and financial development (fdc) is 0.53 and this is relatively close to Tsao's average factor share calculation for physical capital (0.6) for Singapore (Chong-Yah and Lloyd (1986) p. 21). 1/

In contrast to the negative trend in the first cointegrating vector, the trend coefficient in the second cointegrating vector is positive. This corresponds to the evidence in Tsao's (1982) most rigorous work on the sources of growth accounting for the Singapore economy, in which she concluded that total factor productivity for the period 1972-80, though small, is positive. This can be rationalized on the basis of the fact that there is relatively little Research and Development (R & D) expenditure occurring in this economy (*The Economic Survey of Singapore* (1992) p. 63). The positive trend coefficient for the second cointegrating vector is economically meaningful, since one can interpret it as capturing long-run technical progress, and is also in line with Young's most recent (1993) total factor productivity calculations for Singapore with respect to the period 1970-85.

1/ In fact, in his more recent work on a comparative study of Hong Kong and Singapore, Young (1992) estimated a factor share for capital of 0.468 for the period 1985-90 and 0.533 for 1970-90 for Singapore.

Chart 9. The Concentrated Cointegrating Relationships and their Respective Recursively Computed Eigenvalues.



The recursively computed time-series values of the trend coefficient in the second estimated and identified long-run output cointegrating vector can thus be validly be used as a proxy for the long-run total factor productivity in Singapore. The first difference of this variable, that is, the long-run total factor productivity growth rate (Δtfp), is regressed on the first difference of the financial development variable (Δfdc) and real deposit rates of interest (r_o - π), the latter is a stationary I(0) variable. This yields the following regression,

$$\Delta(tfp)_t = \underset{[2.264]}{0.9138} + \underset{[3.707]}{1.006}\Delta fdc_t + \underset{[1.815]}{0.5412}(r_o - \pi)_t$$

$$R^2 = 0.4362 \qquad (8)$$

Thus, in line with the evidence in Fry (1991, p. 31) for 10 Asian developing countries, we find a positive and significant relationship between growth in long-run total factor productivity and financial development in Singapore. However, the credit variable dominates the real deposit rates variable in the regression. Within the context of the long-run augmented production function being analyzed here, this result suggests the efficiency-improving effect of financial development on investment in Singapore (Fry (1991, 1993)).

Moreover, an important question is whether financial development (fdc) and human capital (hc) are necessary variables for cointegration in the aggregate production function VAR. The likelihood ratio test for the null hypothesis that the coefficient on fdc and hc in the first cointegrating vector are both zero yields $\chi^2(2) = 14.9$ compared with its 5 percent critical value of 5.99 and is thus strongly rejected. A similar result was found for the second cointegrating vector. 1/ If either fdc or hc is omitted from the VAR, the Johansen likelihood ratio tests for the null hypothesis of no cointegrating vector cannot be rejected. Thus output, labor, and physical capital are not in themselves cointegrated; both financial development and human capital variables are necessary additional variables for cointegration in Singapore's aggregate production function.

Our two rival long-run equilibrium output equations, which were estimated from the augmented aggregate production function VAR, derived from its cointegrating space, and subsequently interpreted using the new growth theory, are then as follows:

1/ A number of other tests were performed to check the robustness of this result: (i) fdc excluded but hc included; and (ii) hc excluded but fdc included. Neither specification yielded an output cointegrating vector with sensible coefficients.

$$y_t = 0.778k_t + 0.312l_t + 0.412fdc_t + 0.087hc_t - 0.005Trend_t, \quad (9)$$

and

$$y_t = 0.272k_t + 0.473l_t + 0.258fdc_t + 0.113hc_t + 0.002Trend_t, \quad (10)$$

We therefore use equations (9) and (10) to define our error-correction terms ECM1 and ECM2, respectively.

The values of the coefficient of adjustments $\boldsymbol{\alpha}$, in Table 3, indicate the rate of feedback of deviations from the long-run output relationship to the dynamic behavior of the endogenous variables. By inspection from the coefficients in the first column of $\boldsymbol{\alpha}$, it would appear that weak exogeneity is satisfied in the first cointegrating vector. However, if we look at the second column, denoting the feedback coefficients for the second cointegrating vector, it is evident that labor(employment) is not weakly exogenous with respect to the long-run output equation. Given that this is a partial analysis, to maintain rigor, we do a full system test for weak exogeneity in the $\boldsymbol{\alpha}$ loading matrix as formulated in Johansen (1992). 1/ The null hypothesis is H_o: $\alpha_{21} = \alpha_{31} = \alpha_{41} = \alpha_{51} = 0$ and the appropriate matrices are,

$$A = \begin{pmatrix} 1 \\ 0 \\ 0 \\ 0 \\ 0 \end{pmatrix} \qquad B = \begin{pmatrix} 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{pmatrix}$$

The likelihood ratio test for this hypothesis is distributed as χ^2, and this is calculated as 8.87, compared with the critical value of $\chi^2_{0.05}(4) = 9.49$ thus narrowly accepting the weak exogeneity hypothesis in the first cointegrating vector. Repeating a similar exercise for the second cointegrating vector, yields $\chi^2 = 20.32$, compared with the critical value of $\chi^2_{0.05}(4) = 9.49$, thus strongly rejecting weak exogeneity. The latter result is not so surprising, given the cross-equation links evident between output and labor in column 2 of $\boldsymbol{\alpha}$ adjustment coefficients (Table 3). The finding also accords with standard economic theory, for a priori, employment is endogenously linked to output. Consequently, we shall estimate the augmented production function within a bivariate (output, labor) system.

1/ The same linear combinations that are used to identify the error-correction terms from the cointegrating space β', must be appropriately applied to the α loading matrix. However, doing this here does not radically alter the values in the original α loading matrix in Table 3, at least not for the values of α corresponding to the cointegrating space.

IV. A Bivariate System Model of the Aggregate Production Function for Singapore

The cointegration analysis in the aggregate production function has yielded two rival long-run equilibrium output equations, with one implying increasing returns to scale (IRS) and the other, constant returns to scale (CRS). In their theoretical model of financial development and endogenous growth, Berthélemy and Varoudakis (1995) show that there exist multiple steady state equilibria owing to a reciprocal externality between the banking and the real sectors. Multiplicity of equilibria is a basic phenomenon of most general equilibrium systems, but economic theory tells us that often one of the equilibria may be more robust than the others.

Below, we estimate a bivariate system model (output and labor) of the augmented aggregate production function for Singapore. Since the cointegration analysis has yielded two rival long-run output equations, we shall estimate two corresponding bivariate systems with each system augmented by the relevant error-correction term (ECM). Our general bivariate (y, l) system model is of the form:

$$\begin{aligned} A1(L)\Delta y_t = \alpha_{0a} &+ A2(L)\Delta k_t + A3(L)\Delta l_t + A4(L)\Delta fdc_t \\ &+ A5(L)\Delta hc_t + A6(L)\Delta fdi_t + A7(L)\Delta \pi_t \\ &+ A8(L)ECM_{t-1} + A9TREND_t + \epsilon_{ta}, \end{aligned} \tag{11}$$

and

$$\begin{aligned} B1(L)\Delta l_t = \alpha_{0b} &+ B2(L)\Delta k_t + B3(L)\Delta y_t + B4(L)\Delta fdc_t \\ &+ B5(L)\Delta hc_t + B6(L)\Delta fdi_t + B7(L)\Delta \pi_t \\ &+ B8(L)ECM_{t-1} + B9TREND_t + \epsilon_{tb}, \end{aligned} \tag{12}$$

where A1(L).....A8(L) and B1(L).....B8(L) are polynomials of the form $A(L) = \sum\alpha_r L^r$ and $B(L) = \sum\alpha_r L^r$, respectively, in which L is a lag operator, such that $L^i x_t = x_{t-1}$. Foreign direct investment (fdi_t) and inflation π_t are taken as exogenously given in the production function model, α_{0i} are constants, and ϵ_{ti} are the error terms. Throughout the reduction sequence various parameterizations of the variables were considered for each system equation, but the specifications adopted here variance dominate all the other alternative specifications. Performing the estimation in the augmented aggregate production function using a Recursive Full Information Maximum Likelihood (RFIML) estimator, results in the following parsimonious representations of the bivariate system:

System 1. RFIML Bivariate Model of Output and Labor

$$\begin{aligned}\Delta y_t = {} & \underset{[1.382]}{0.007} + \underset{[2.719]}{0.259\Delta y_{t-1}} + \underset{[3.152]}{0.402\Delta fdc_{t-2}} + \underset{[3.878]}{0.704\Delta k_{t-1}} \\ & + \underset{[2.377]}{0.883\Delta l_{t-1}} + \underset{[1.855]}{0.496\Delta hc_{t-3}} + \underset{[2.984]}{1.637\Delta fdi_{t-1}} \\ & - \underset{[-4.322]}{0.076 ECM1_{t-1}}\end{aligned}$$

$$\text{Equation } \sigma = 0.0206 \qquad (13)$$

$$\begin{aligned}\Delta l_t = {} & \underset{[-0.791]}{-0.0012} + \underset{[4.832]}{0.595\Delta l_{t-1}} + \underset{[3.069]}{0.107\Delta y_{t-1}} + \underset{[2.614]}{0.092\Delta fdc_{t-2}} \\ & - \underset{[-2.407]}{1.203\Delta (k/l)_{t-2}} + \underset{[2.243]}{0.179\Delta fdi_{t-3}} - \underset{[-2.185]}{0.002 TREND_t} \\ & - \underset{[-1.906]}{0.0418 ECM1_{t-1}}\end{aligned}$$

$$\text{Equation } \sigma = 0.0037 \qquad (14)$$

System 2. RFIML Bivariate Model of Output and Labor

$$\begin{aligned}\Delta y_t = {} & \underset{[0.980]}{0.005} + \underset{[2.246]}{0.214\Delta y_{t-1}} + \underset{[2.993]}{0.382\Delta fdc_{t-2}} + \underset{[3.687]}{0.684\Delta k_{t-1}} \\ & + \underset{[2.314]}{0.857\Delta l_{t-1}} + \underset{[1.796]}{0.479\Delta hc_{t-3}} + \underset{[2.851]}{1.568\Delta fdi_{t-1}} \\ & - \underset{[-5.072]}{0.060 ECM2_{t-1}}\end{aligned}$$

$$\text{Equation } \sigma = 0.0174 \qquad (15)$$

$$\begin{aligned}\Delta l_t = {} & \underset{[-0.468]}{-0.0019} + \underset{[3.767]}{0.573\Delta l_{t-1}} + \underset{[2.598]}{0.094\Delta y_{t-1}} + \underset{[2.493]}{0.0862\Delta fdc_{t-2}} \\ & - \underset{[-2.552]}{0.984\Delta (k/l)_{t-2}} + \underset{[2.084]}{0.155\Delta fdi_{t-3}} - \underset{[-2.091]}{0.001 TREND_t} \\ & - \underset{[-3.607]}{0.0707 ECM2_{t-1}}\end{aligned}$$

$$\text{Equation } \sigma = 0.0031 \qquad (16)$$

System Model 1 denotes the bivariate (y,l) system representation of the augmented aggregate production function in which the increasing returns to scale long-run output equation is the error-correction term (ECM1). Similarly, Model 2 represents the bivariate system (y,l) in which the constant returns to scale (CRS) long-run output equation is the error-correction term (ECM2). Both systems consist of estimated parsimonious first difference models of output and labor augmented by the relevant ECM cointegrating vector.

It is evident from both systems, that the financial development variable enters significantly into both the growth and the labor equations. However, both human capital (hc) and inflation (π) turn out to be insignificant in the estimated bivariate system models. The significance of Δy_{t-1} (one-period lagged output) in the labor equation echoes the result of the earlier weak exogeneity test, but it can also have an interesting economic interpretation. Fluctuations in expected output are important factors in explaining employment fluctuations in the short run (Nickell (1984) and Harvey et al (1986)). Thus Δy_{t-1} can be interpreted as incorporating intertemporal optimization subject to adaptive expectations in output. The significance of the capital-labor ratio (k/l) in the labor equation reflects the effect of relative factor prices on labor input. Foreign direct investment (fdi) enters the labor equation in both bivariate systems, albeit with a three-lagged period. A bulk of investment in Singapore's manufacturing sector consists of direct investment from overseas and since manufacturing and financial sector are the two largest employers in Singapore, this is a plausible finding. The coefficient on the TREND term in the labor equation implies approximately 0.1 percent to 0.2 percent reduction in employment per quarter, owing to technical progress. The parameters of the labor equation for Singapore, therefore, seem quite sensible.

The analysis of the estimated labor functions in this section of the paper should not be considered as a digression, for it is the direct result of the absence of weak exogeneity in the estimated augmented aggregate production function. This necessitated the system estimation technique employed above. Nonetheless, the key theme addressed in this paper is growth. Consequently, the rest of the paper will focus on this issue. The estimated growth models from the two systems will serve as a laboratory to further test stringently the robustness of the two rival cointegrated long-run solutions for output. Below we report the two rival growth models from the respective systems along with their diagnostics statistics.

$$\Delta y_t = \underset{[1.382]}{0.007} + \underset{[2.719]}{0.259\Delta y_{t-1}} + \underset{[3.152]}{0.402\Delta fdc_{t-2}} + \underset{[3.878]}{0.704\Delta k_{t-1}} + \underset{[2.377]}{0.883\Delta l_{t-1}} + \underset{[1.855]}{0.496\Delta hc_{t-3}} + \underset{[2.984]}{1.637\Delta fdi_{t-1}} - \underset{[-5.472]}{0.076 ECM1_{t-1}}$$

T = 1971(1) - 1991(2) *less 10 forecasts,*

R^2 = 0.70, σ = 0.0206, DW = 1.87,

Chow $F(10,64)$ = 0.36, *Forecast* $\chi^2(10)/10$ = 0.39,

Normality $\chi^2(2)$ = 0.646, *AR* 1-4 $F(4,60)$ = 1.15,

ARCH $F(4,56)$ = 0.58, X_iX_j $F(14,49)$ = 0.67, *RESET* $F(1,63)$ = 0.17, (17)

$$\Delta y_t = \underset{[0.987]}{0.005} + \underset{[2.247]}{0.214\Delta y_{t-1}} + \underset{[2.997]}{0.382\Delta fdc_{t-2}} + \underset{[3.768]}{0.684\Delta k_{t-1}} + \underset{[2.307]}{0.857\Delta l_{t-1}} + \underset{[1.791]}{0.479\Delta hc_{t-3}} + \underset{[2.858]}{1.568\Delta fdi_{t-1}} - \underset{[-4.322]}{0.060 ECM2_{t-1}}$$

T = 1971(1) - 1991(2) *less 10 forecasts,*

R^2 = 0.79, σ = 0.0174, DW = 2.03,

Chow $F(10,64)$ = 0.26, *Forecast* $\chi^2(10)/10$ = 0.28,

Normality $\chi^2(2)$ = 0.585, *AR* 1-4 $F(4,60)$ =1.01,

ARCH $F(4,56)$ = 0.46, X_iX_j $F(14,49)$ = 0.59, *RESET* $F(1,63)$ = 0.10, (18)

Table 4. Competing Growth Model Performance

Growth Model Equation	Equation Std Error	Schwarz Criterion 1/	Forecast $\chi^2(10)/10$
Eq (17)	0.0206	-7.91	0.39
Eq (18)	0.0174	-7.26	0.28

Table 5. Non-Nested Encompassing Tests
(Growth Model 17 versus 18)

Model(17) vs Model(18)	Form	Test	Form	Model(18) vs Model(17)
-0.956	N(0,1)	Cox	N(0,1)	0.0797
-0.900	N(0,1)	Ericsson IV	N(0,1)	-0.0752
0.0677	$\chi^2(1)$	Sargan	$\chi^2(1)$	0.0106
0.0667 (0.7970)	F(1, 62)	Joint	F(1, 62)	0.0104 (0.9189)

1/ The Schwarz statistic is a measure of the likely forecast efficiency of a model, Charemza and Deadman (1992). It is defined as follows; SC = $\ln\sigma^2 + k\ln T/T$, where σ^2 is a degree-of-freedom adjusted equation standard error (squared), k the number of parameters and T the sample size. SC is thus increasing in σ^2 and in k: a fall in the value of SC is an indication of undominated parsimony.

The identification of any robust econometric specification involves a series of iterative processes across the variables and their dynamics. Thus we commence by assessing the competing growth models (estimated over the same data and sample period) relative to an independent performance criterion. The relative performance of the models is compared in terms of basic diagnostic statistics and non-nested encompassing tests respectively in Table 4 and Table 5 above. Table 4 indicates that in terms of basic battery of diagnostic tests required in a model design, the CRS growth equation model (18) generally outperforms the IRS growth equation model (17), most notably in terms of forecast accuracy.

However, the diagnostic statistics do not answer the question as to which of the models is better in a direct comparison with each other, that is whether one of them should be rejected in favor of the other. In principle, such a selection can be made by the formulation and testing of non-nested hypotheses. Table 5 above summarizes the results of the non-nested encompassing tests. Under the null that Model 1 encompasses Model 2, both the Cox test and the Ericsson IV test have a standard normal distribution. The Sargan test has a χ^2 distribution and the F test is a test of each model against the joint nesting model. Details of the individual tests are provided in Hendry and Doornik (1993, pp. 19-22). From the table, it is evident that the growth equation model (18) significantly encompasses the growth equation model (17), which therefore reinforces the results obtained from the comparative diagnostic statistics in Table 4. Consequently, the second long-run solution (ECM2) that was derived by simple row reduction in the cointegrating matrix space and which satisfies constant returns to scale (CRS), encompasses the increasing returns to scale (IRS) cointegrating vector (ECM1). The CRS vector thus dominates the IRS vector in terms of being the disequilibrium feedback term in the estimated error-correction dynamic growth model for the Singapore economy. 1/

The results above are not however consistent with the limited comparable evidence on Singapore's aggregate production function. As cited above, Chen (1993) and Kim and Lau (1993) rejected the Cobb-Douglas specification within a translog production function model developed by Christensen, Jorgenson and Lau (1973). Similarly, using standard neoclassical growth accounting techniques, Young (1992) computed total factor productivity for Singapore from a translog production function and found this to be negative for a large part of the sample period (1974-89). Leaving for the moment the degree of rigor required for estimating a long-run translog production function (Granger and Hallman (1991)) on the unit-roots and cointegrated properties of non linear

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1/ A log-likelihood ratio test for over-identifying restrictions between the estimated CRS growth model in equation 18 and the more general growth model equation 11 yielded, $\chi^2(33) = 19.86$. Thus, the estimated CRS growth model in equation 18 encompasses not only the IRS growth model in equation 17 but also the more general growth model in the bivariate system in equation 11.

transformations of integrated time series), 1/ the explanation for these results can be twofold. First, by ignoring both the human capital (hc) and the financial development (fdc) variables, Young (1992), Chen (1993) and Kim and Lau (1993) result could be due to omitted variables bias. This could be driving the huge elasticity values they found for $(k)^2$, $(1)^2$ and (1.k) which not only led to a rejection of the Cobb-Douglas specification, but also gave rise to a perverse trend coefficient, implying negative technical progress. The latter result is observationally equivalent to a negative total factor productivity, but in contrast, our robust constant returns to scale (CRS) cointegrating vector has yielded a positive technical progress coefficient for Singapore. Therefore, the negative technical progress coefficient embodied in ECM1, and found by Young (1992) 2/ and Kim and Lau (1993) is not a robust result for the dynamic long-run growth path of the Singapore economy. 3/

In summary then, the Cobb-Douglas specification for Singapore's aggregate production function, in which the basic factor inputs of physical capital and labor are augmented by human capital and financial development variables, has yielded a CRS long-run output cointegrating vector that is consistent with the data, explains previous findings and thus encompasses them. The CRS long-run output equation is found to be an adequate disequilibrium feedback term in the estimated parsimonious growth model and provides a basis for a coherent economic interpretation for the underlying growth model for Singapore.

1/ Granger and Hallman (1991) have shown that once one has squared terms of integrated series, one moves from the Dickey-Fuller (1981) territory to a different asymptotic theory. They recommended the Ranked Dickey-Fuller test and a different cointegration theory.

2/ In his more recent work on East-Asian NICs, in which he did not employ the translog production function model, Young (1993) found positive total factor productivity time series values for Singapore.

3/ Since three of the variables in the augmented aggregate production function, namely physical capital, human capital and labor were interpolated into quarterly series in order to carry out estimation, sensitivity analysis was done so as to evaluate the robustness of the estimated cointegration results using the annual series. The estimated long-run solutions from the annual data were found to be quite comparable to the quarterly ones reported here.

V. A Production Function Growth Model for Singapore: Statistical Robustness and Economic Interpretation

Model equation (18) is the dominant production function growth model for Singapore. Having encompassed the rival IRS growth model, we now evaluate it further by examining the statistical robustness of the estimated growth model. This is achieved here by subjecting the dominant ECM growth model from the bivariate system to a general battery of tests involved in a model design, namely: model stability, one-step equation residuals, parameter constancy of the significant variables, and the forecasting performance of the estimated growth model.

The overall model can be gauged from Chart 10 (the one-step Chow test) and Chart 11 (the recursive equation residuals) and these indicate a high degree of model stability. A more comprehensive analysis of the parameter stability of the model can be achieved by using a recursive system (RFIML) estimator. Because of the degrees-of-freedom limitation on the recursive system estimators, it is not possible to examine the recursive behaviour of the coefficients before 1975, which is unfortunate in view of the fact that a major part of the financial liberalization program started during this period and thus any noticeable parameter instability (due inter alia to the absence of super exogeneity) may be expected to emerge. Nonetheless, the recursive system coefficient plots shown in Charts 12 - 17 indicate a similar picture of parameter and model stability. In particular, the error-correction term ($ECM2_{t-1}$) shows a high degree of parameter constancy. All the coefficients of the significant regressors of the growth model are well inside the ex ante standard errors, with the standard error bounds converging over time. The next pages report the graphs of model and parameter stability of Singapore's dominant growth equation model.

Chart 10. One-Step System Chow Test for a Model of Δy_t, with Statistics Scaled by their One off 5 Percent Critical Values.

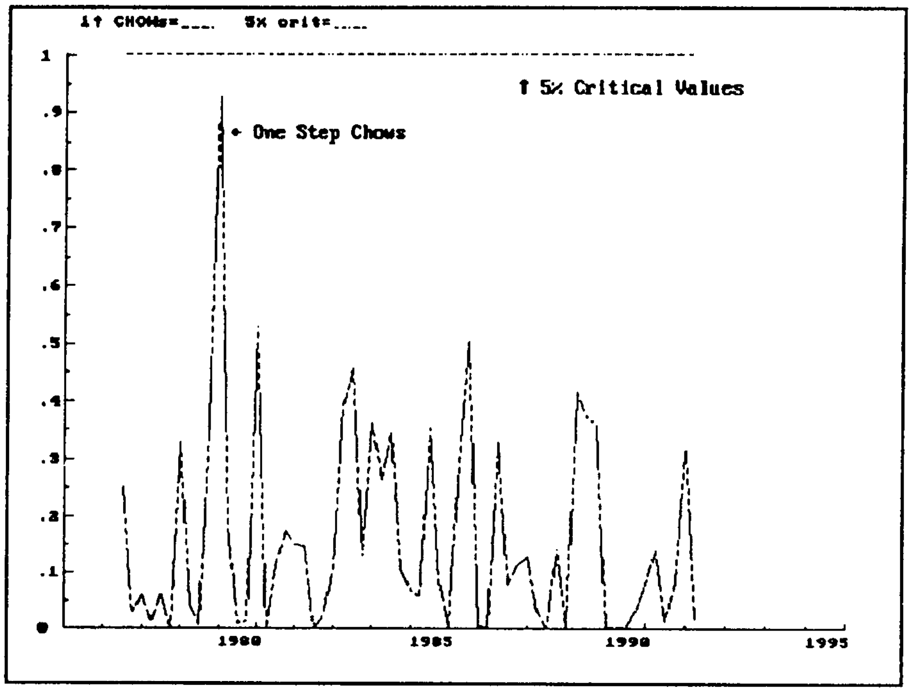


Chart 11. One-Step System Residuals and the Corresponding Calculated Equation Standard Errors (s.e.) for a Model of Δy_t.

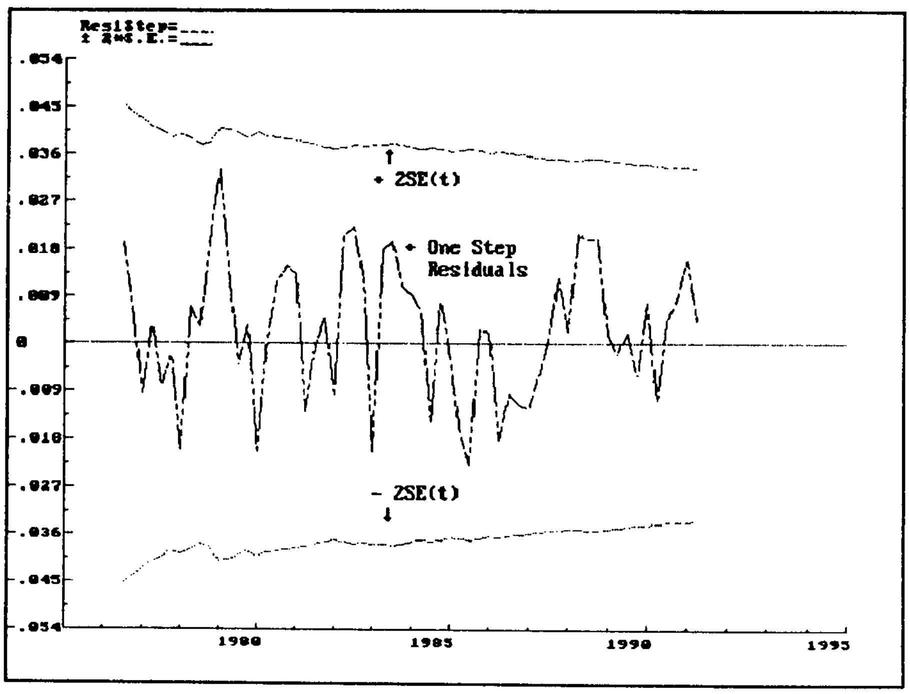


Chart 12. Recursive System (RFIML) Estimates of the Coefficient of Δy_{t-1} for the Model Δy_t, with ±2 Estimated Standard Errors.

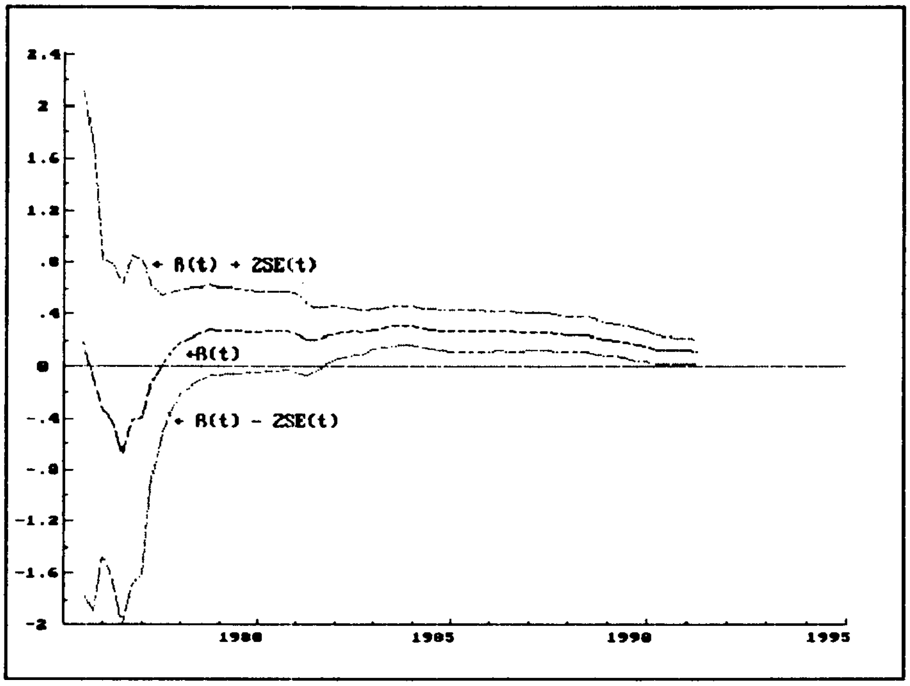


Chart 13. Recursive System (RFIML) Estimates of the Coefficient of Δfdc_{t-2} for the Model Δy_t, with ±2 Estimated Standard Errors.

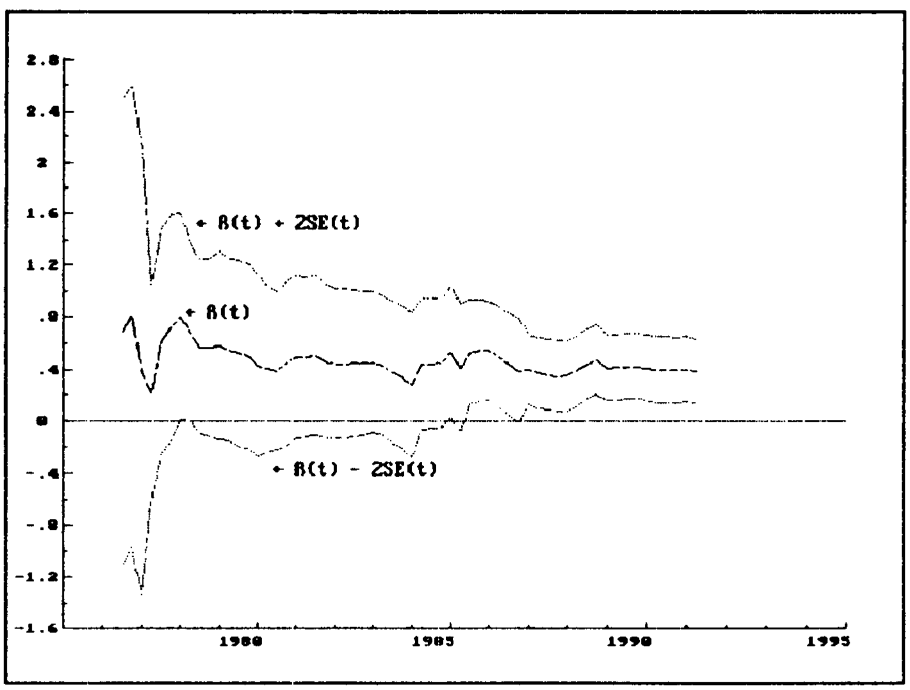


Chart 14. Recursive System (RFIML) Estimates of the Coefficient of Δk_{t-1} for the Model Δy_t, with ±2 Estimated Standard Errors.

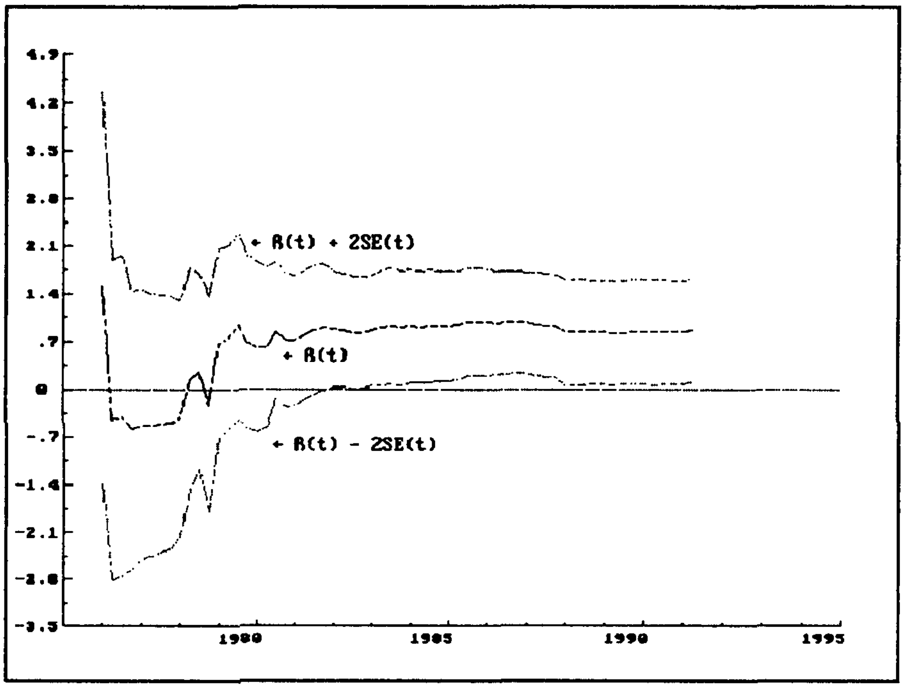


Chart 15. Recursive System (RFIML) Estimates of the Coefficient of Δl_{t-1} for the Model Δy_t, with ±2 Estimated Standard Errors.

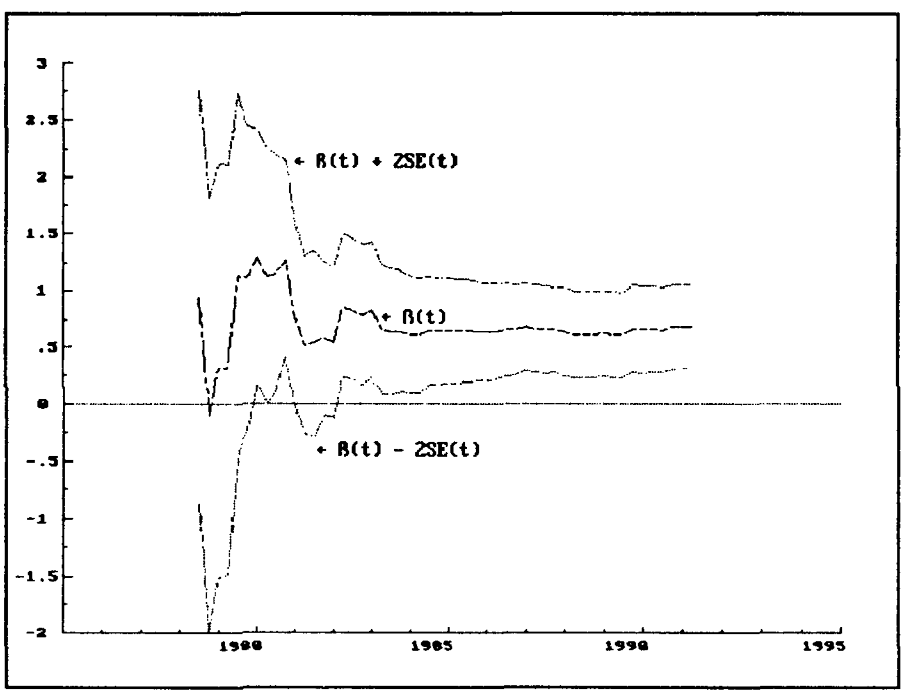


Chart 16. Recursive System (RFIML) Estimates of the Coefficient of Δfdi_{t-1} for the Model Δy_t, with ±2 Estimated Standard Errors.

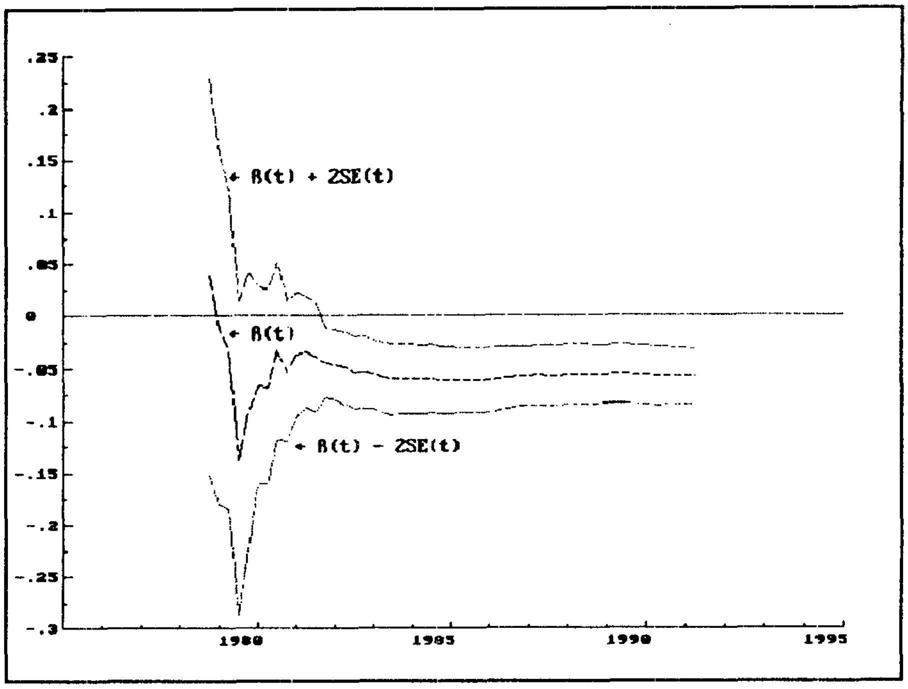
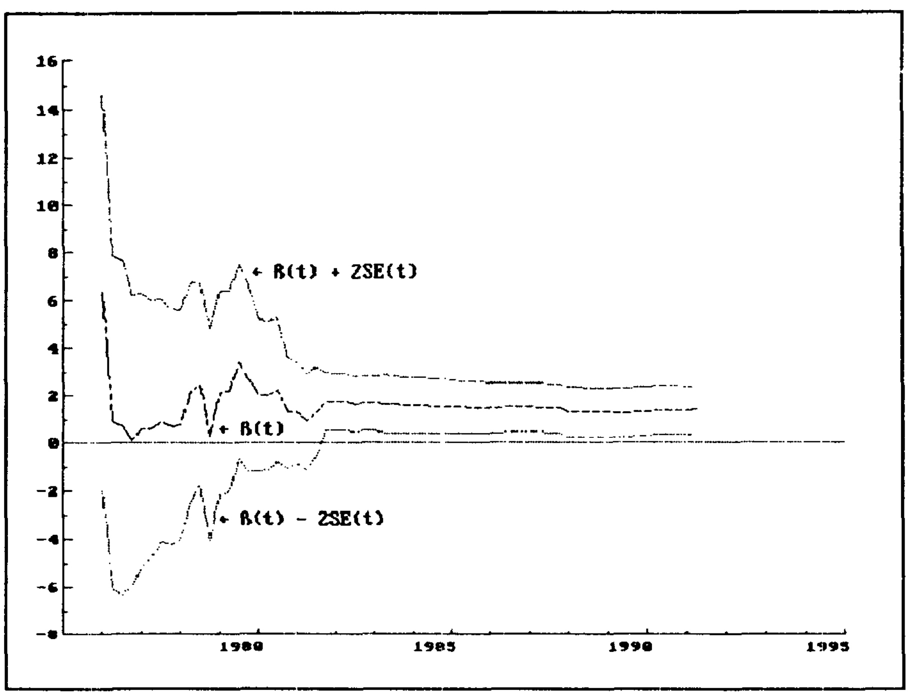


Chart 17. Recursive System (RFIML) Estimates of the Coefficient of ECM_{t-1} for the Model Δy_t, with ±2 Estimated Standard Errors.



In the derivation of the recursive system (RFIML) coefficient values of the significant variables of the growth model equation (18), different degrees of initialization are used in order to maximize on intertemporal efficiency. In each case, the initialization chosen for each variable yielded the optimal degree of parameter constancy over time. Finally, the direct test for forecast stability of the model confirms that this cannot be rejected for the last ten quarters and the graph of actual and fitted values of Δy_t in Chart 18 indicates that the latter tracks the former well throughout the sample period. The plot of actual and forecast values of Δy_t, shown in the same Chart illustrates good model forecast performance. Table 6 reports the statistics of the forecast analysis for the growth model equation (18).

The growth model equation (18) has satisfied the basic classical regression conditions of error homoscedasticity or zero autocorrelation and passed the standard battery of diagnostic tests involved in a model design. Extra statistical scrutiny of the model, through the recursive system full information maximum likelihood (RFIML) analysis, indicates the robustness of various parameter system estimates of the growth model. The short-run dynamic processes of the growth model are felt within a relatively short period; in no case are there significant lagged effects beyond two quarters. Thus the process of reduction has increased efficiency over the nesting model equation (11) and allowed for a clearer specification of the dynamic processes of the augmented aggregate production function system.

In terms of economic interpretation, growth model equation (18) indicates that transitional growth in Singapore is determined by financial development, physical capital stock, labor input, and foreign direct investment. As evident from this growth model, human capital does not significantly affect Δy_t except through the error-correction term (ECM_{t-1}). Thus, although human capital affects growth in the long run, as evident in the two estimated cointegrating vectors, it does not affect short-run or transitional growth. Inflation (π) enters insignificantly in the growth model. However, the labor input variable significantly affects growth both in the short run and the long run. This implies that any policy that reduces the natural rate of unemployment, thereby increasing labor input, will potentially boost output growth in Singapore.

The financial development variable, proxied by credit influences growth both in the short run as shown by the high level of significance of Δfdc_{t-1} in the growth equation (18), as well as in the long run through the error-correction term ($ECM2_{t-1}$). Note that there is a slight deviation from trend of the recursively computed coefficients $\beta(t)$ for Δfdc_{t-1} from 1984 to 1986 (Chart 12). This partly suggests that credit was probably an important factor in the 1985 downturn, the first major recession experienced by the Singapore economy over the post war era. This, combined with the interaction identified between the financial development variable (fdc) and the physical capital stock variable (k) in the dominant estimated long-run output equation has a logical economic interpretation. Financial development in Singapore influences short-run or transitional growth, but in the long run, it affects growth through its positive impact on investment efficiency.

Chart 18. Actual, Fitted and Forecast Values of the Estimated Growth Model Δy_t and Forecast over the Period 1989Q2-1991Q2.

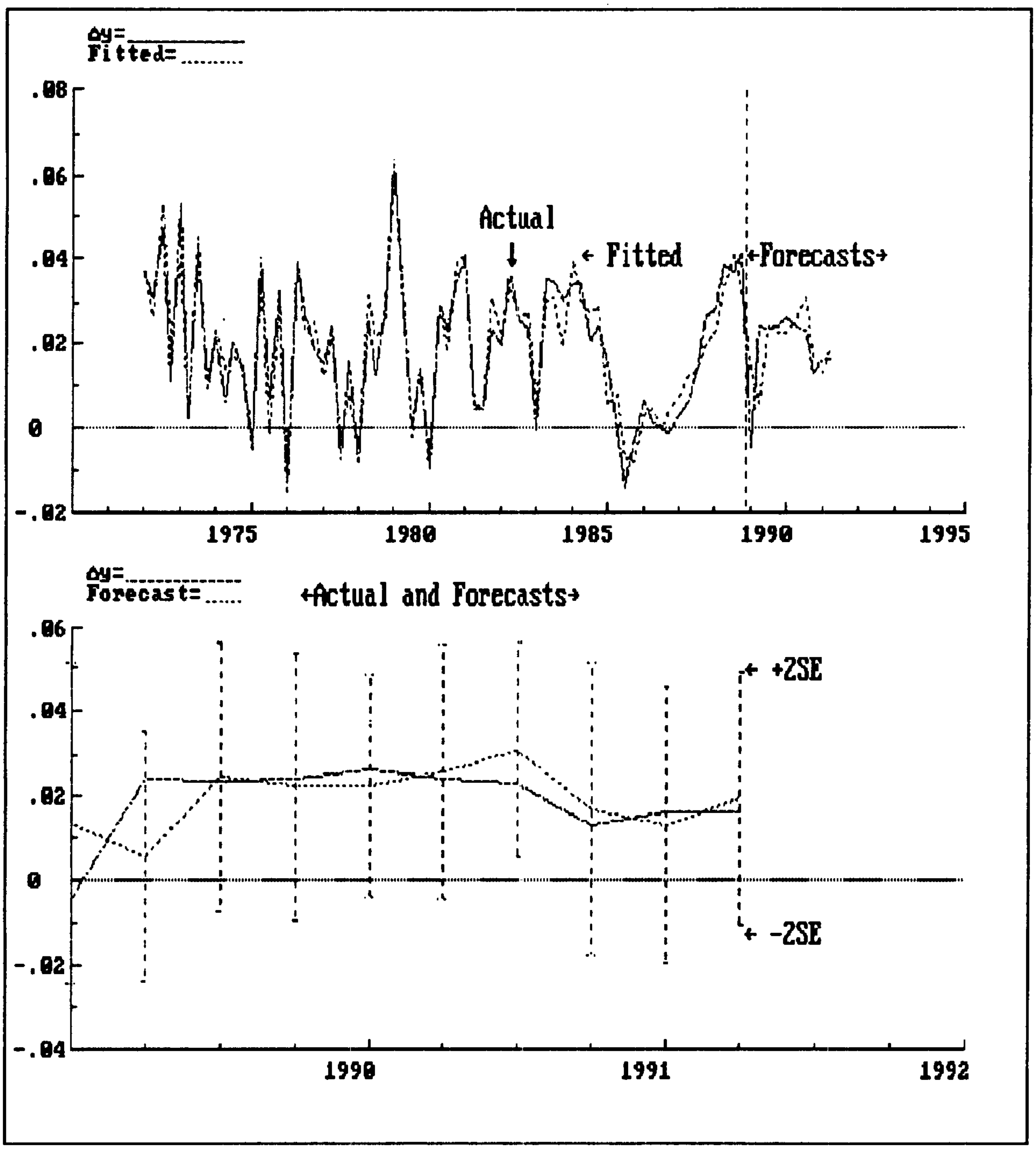


Table 6. One-Step Forecast Analysis for Growth Model Equation (18) and Forecast over the Period 1989Q1-1991Q2

Date	Actual(Δy)	Forecast(Δyhat)	Δy - Δyhat	Forecast SE	t-Value
1989Q1	-0.00408	0.01362	-0.01770	0.01878	-0.9424
1989Q2	0.02405	0.00589	0.01816	0.01483	1.2243
1989Q3	0.02352	0.02473	-0.00121	0.01595	-0.0758
1989Q4	0.02411	0.02235	0.00176	0.01569	0.1127
1990Q1	0.02638	0.02240	0.00398	0.01307	0.3045
1990Q2	0.02397	0.02593	-0.00196	0.01501	-0.1305
1990Q3	0.02290	0.03105	-0.00815	0.01268	-0.6428
1990Q4	0.01316	0.01694	-0.00378	0.01732	-0.2183
1991Q1	0.01660	0.01331	0.00329	0.01619	0.2031
1991Q2	0.01646	0.01970	-0.00324	0.01485	-0.2178

Tests of Parameter Constancy over: 1989Q1 - 1991Q2

Forecast $\chi^2(10)/10$ = 0.2794

Chow Test F(10, 64) = 0.2595 (0.9876)

Zero Forecast Innovation t(9) = 0.0956

In the traditional literature of McKinnon-Shaw (1973) and Kapur-Mathieson (1981, 1980a), it is largely through the effect of positive real interest rates on savings, and hence investment, that financial development enhances economic growth. However, the new endogenous growth models argue that the relationship between interest rates, savings, and investment is murky (De Gregorio (1992a) and Fry (1993)) and that is a weakness of the McKinnon-Shaw (1973). The dominant error-correction growth model estimated from the augmented aggregate production function by means of systems modeling in this paper strongly suggests that the main channel of the transmission mechanism from financial development to growth in Singapore, in the medium to long run, is through its positive effect on the efficiency of investment. Moreover, using the growth equation model (18), average per quarter growth in Singapore was decomposed into endogenous and exogenous factors. Taking note of the fact that the constant term α_{0i} and foreign direct investment (fdi) are the exogenous variables of the estimated growth model (18), while the rest of the variables are the endogenous variables of the Johansen system, the decomposition reveals that 80 percent of growth in Singapore is due to endogenous factors compared with only 20 percent of exogenous growth.

In line with the evidence in Fry (1993), according to our estimated model, foreign direct investment is a significant determinant of growth in Singapore. The development of Singapore into an international financial center provided a further explanation for the flow of foreign direct investment into Singapore. Other factors include political and social stability, good physical and social infrastructure and a disciplined and skilled workforce. However, the effect of fdi on the long-run growth path of the Singapore economy is an important issue that has not been examined in the above econometric analysis. This is evident from the absence of foreign direct investment in the estimated Johansen type cointegrated solutions for output analyzed in this paper. 1/ The implicit exogeneity assumption on fdi in our model, and its consequent absence in our Johansen "closed system" cointegration results, is not entirely satisfactory in view of the observations made above. More generally, foreign direct investment in developing economies is largely influenced by domestic macroeconomic conditions.

Notwithstanding this drawback, our error-correction growth model for Singapore, estimated by system modeling in the augmented aggregate production function, has yielded parameter estimates that have a number of strong features. The dynamic growth model has passed the general battery of

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1/ Part of the problem here, though, is due to the limitation of the current version of PC-FIML (version 7.01-1994), which can only estimate cointegrating relationship(s) in a "closed system" (that is, with all variables being endogenous) and not an "open system" where the long-run cointegrating solution is a function of both endogenous and exogenous variables of the system. PC-FIML version 7.01-1994, does not have facilities for incorporating a stochastic trend in the long-run solution nor the means of including exogenous variables in the cointegrated space.

diagnostic tests involved in a model design and is statistically robust; it is economically interpretable and more importantly, the model has identified both short-run growth factors in Singapore (financial development, inter alia) and long-run determinants. Furthermore, the model has picked up effects that reflect developments on the macroeconomic scene in Singapore and the growth-enhancement policies that have been pursued over the years. These, combined with the within-sample forecast performance and the evidence from the encompassing tests allow us to conclude that the growth model (18) provides a tentatively adequate characterization of the dynamics of growth in Singapore.

VI. Conclusions

Recent cointegration techniques, which focus on the estimation and the identification of long-run economic relationships between data variables are particularly appropriate to the study of long-run endogenous growth models. This paper has applied these techniques to the Singapore data, using a supply-side framework. By and large, the econometric analysis in this paper has yielded results that are in line with predictions of endogenous growth models.

First financial development as proxied here is found to be a significant determinant of the long-run output path of the Singapore economy. Second the human capital variable in Singapore is found to play a significant role in long-run cointegrated solutions for output and this supports another endogenous growth hypothesis. Furthermore, the dominant error-correction growth model estimated in a bivariate system suggests that a significant part of growth in Singapore is driven by endogenously determined variables with a relatively lower contribution from exogenous factors. Moreover, the ECM growth model shows that financial development enhances growth in the long run, mainly through the positive effect on the efficiency of investment. This latter finding is also in line with predictions of the new endogenous growth models.

The estimated augmented aggregate production function has yielded a cointegrating vector for output that exhibits increasing returns to scale (IRS) and another cointegrating vector that satisfies a constant returns to scale (CRS) model. Although some endogenous growth models assume CRS (Fry (1993) and Berthélemy and Varoudakis (1995)), to the extent that the constant returns to scale (CRS) long-run solution for output encompasses the increasing returns to scale (IRS) cointegrating vector in the error-correction growth model, the Singapore data have exhibited characteristics not strictly consistent with all endogenous growth models. Furthermore, the dominant ECM growth model indicates that although human capital influences growth in Singapore in the long run, it does not affect it in the short run. Contrary to previous evidence on negative total factor productivity for this economy, we find a positive coefficient on the trend term in the dominant long-run cointegrated equation for output, which suggests that technical progress in Singapore, though relatively small is positive.

The techniques used and the growth model results in this paper are not definitive. Our method of capturing long-run technical progress in the cointegrated space employs a deterministic trend, which presupposes that technical progress is time-invariant. However, there are good a priori reasons why technical progress may be stochastic in nature. Moreover, the absence of foreign direct investment in our Johansen long-run cointegrated solutions for output is a weakness of the growth models here. Software limitations abound, and therefore we have not been able to address these underlying weaknesses of the models.

Notwithstanding these limitations, this paper has successfully established robust econometric models of endogenous growth for Singapore that have shown the importance of financial development in both the short and the long run. Moreover, the estimated models reflect policies that have been pursued by the Singapore government over the years to enhance growth. Consequently, these enable us to accept that the estimated models in this paper represent an adequate characterization of the factors determining growth in Singapore.

Finally as mentioned already, there is evidence that financial development enhances long-run growth in Singapore partly by improving the efficiency of investment. This suggests that in fulfilling this function the efficiency of asset-pricing in financial markets in Singapore becomes an issue of some importance. Therefore, one possible extension of this paper is an examination of both the short-run and long-run efficiency characteristics of financial markets in Singapore.

APPENDIX I

The Johansen-Juselius (1990) Cointegration Methodology

Following Johansen and Juselius (1990), consider a VAR of the form,

$$X_t = \Pi_1 X_{t-1} + \dots\dots + \Pi_k X_{t-k} + \mu + \Theta D_t + \epsilon_t \quad (t = 1, \dots T), \tag{1}$$

where X_t is a sequence of I(1) vectors with components $(X_{1t}, \dots\dots X_{pt})$, μ is a vector of constants, D is a vector of centered seasonal dummies, and the innovations to this process $\epsilon_1 \dots\dots \epsilon_T$, are IN(0, Ω). Using the difference operator, equation (1) can be reparameterized as,

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \dots\dots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + \Theta D_t + \epsilon_t, \tag{2}$$

where $\Gamma = -(I - \Pi_1 - \dots - \Pi_i)$ $(i = 1 \dots k - 1)$
and $\Pi = -(I - \Pi_1 - \dots\dots - \Pi_k)$.

Equation (2) combines both short-run and long-run dynamics and it is the matrix Π that conveys information about the long-run relationship between the X variables.

The X_t vector processes are assumed to be integrated of order one-I(1), but the number of cointegrating vectors v, and hence the number of unit-roots N - v, is unknown. The distributions of statistics are non-standard in such a setting and require special critical values. Testing the size of v is equivalent to testing whether $\Pi = -\alpha\beta'$, where β and α are N $\times$ v matrices. Since the likelihood function depends on the normal distribution, it can be concentrated with respect to Ω, ϕ and $(\Pi_1 - \dots\dots - \Pi_{t-k})$, the last by obtaining residuals R_{ot} and R_{pt} from regressing ΔX_t and X_{t-p}, respectively, on $\{\Delta X_{t-i}\}$. Denote the second moment matrices from these residuals by S_{oo}, S_{op}, S_{po}, and S_{pp}, where $S_{ij} = T^{-1}\Sigma R_{it}R_{jt}'$ for i, j = 0 and p. Then v is determined by the largest eigenvalues $\mu_1 \geq \dots\dots \geq \mu_v \dots\dots \geq \mu_N$ of

$$\left| \mu S_{pp} - S_{po} S^{-1}_{oo} S_{op} \right| = 0, \tag{3}$$

and β' by the corresponding eigenvectors. The maximized likelihood is given by

$$L(\beta') = A - 1/2\ T \textstyle\sum^{v}_{i=1} \ln(1 - \mu_i), \tag{4}$$

and tests of the hypothesis that there are $0 \leq v \leq N$ cointegrating vectors can be based on

$$\text{Trace} = -T \sum^{v}_{i=1} \ln(1 - \mu_i), \tag{5}$$

(twice the logarithm of the likelihood ratio for restricting Π), with v being selected through the first significant statistic of the trace. Alternatively, sequential tests of significance of the largest (μ_v) can be based on the *maximal eigenvalue* that is given by $-T \ln (1 - \mu_v)$. Under the null hypothesis that the eigenvalues are zero, both the *trace* and the *maximal eigenvalue* have distributions that are functionals of Brownian motion, but the critical values for these tests can be found in Johansen and Juselius (1990) and Osterwald-Lenum (1990), among others.

Cointegration may be detected by examining the Π matrix. The number of cointegrating vectors, v, between the elements of X, will therefore determine the rank of the matrix Π. There are three alternative situations, depending on the rank of Π. If Π is of full rank then the matrix is stationary, and the data are not I(1). If the rank is zero, then the variables are all individually I(1) but not cointegrated, in which case an error-correction (ECM) isomorphism cannot be employed. If the rank is greater than zero but less than p, there are v cointegrated vectors, which can be identified and embedded in an error-correction model (ECM). When $0 < v < p$, Π can be factorized as $\Pi = -\alpha\beta'$, where α and β are $v \times p$, with β' containing the v cointegrating vectors and α their corresponding adjustment or feedback coefficients. The matrix β' represents the parameters of economic interest in the VAR and its vectors constitute the long-run (levels) relationships between the variables. The cointegration combinations derived from the eigenmatrix β' form the basis of defining the error-correction model (ECM). Moreover, α reveals the importance of each cointegrating combination in each equation of the VAR. If a given ECM enters more than one equation, then the parameters are cross-linked between such equations, violating weak exogeneity and entailing joint estimation for efficient inference (Hendry and Mizon (1990, 1993)).

Appendix II

Data Definitions and Sources

Data Variable

1. GNP(Gross National Product - Expenditure Based).
   Source: International Department, South East-Asia Division of the Bank of England.1/

2. Credit from Financial Institutions to the Private Sector.
   Source: IFS (International Financial Statistics) of the IMF.

3. Foreign Direct Investment
   Source: World Bank, STARS Database, Washington, D.C.

4. CPI (Consumer Price Index), 1985 = 100.
   Source: IFS (International Financial Statistics) of the IMF.

5. Physical Capital Stock.
   Source: The International Economics Department, The World Bank, STARS Database.2/

6. Human Capital Stock.
   Source: The International Economics Department, The World Bank, STARS Database.

7. Labor-Hours (Employment and Hours of Work).
   Source: Singapore Year book of Statistics, Various Years.

Note: Due to lack of quarterly data, data no. 3, 5, 6 and 7 have been interpolated. While there is no universally accepted method for generating data series through interpolation, this paper has employed Maddala's (1977, pp. 205-207) interpolation method, a variant of Friedman's (1962) procedure. By design the interpolated series are thus consistent with known annual observations.

1/ Thanks to Paul Brockell at Section 9 of the International Division of the Bank.
2/ I am grateful to Vikram Nehru for supplying the data; he is one of the authors of the New Database on Physical and Human Capital Stock at the International Economics Department of the Bank.

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