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SOURCES OF GROWTH REVISITED: THE IMPORTANCE OF THE NATURE OF TECHNOLOGICAL PROGRESS

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Traditional sources of growth studies generally assume that the nature of technological progress is Hicks-neutral. However, the nature of technological progress compatible with steady state conditions is Harrod-neutral rather than Hicks-neutral. This study thus investigates sources of growth for Hong-Kong, the Republic of Korea, Singapore and Taiwan using the bounds testing procedure of Pesaran, Shin and Smith (2001) and the autoregressive distributed lag (ARDL) approach of Pesaran and Shin (1999). The robustness of the test results and parameter estimates are also justified by the fully modified ordinary least squares approach of Phillips and Hansen (1990). The results emphasize that the fundamental source of economic growth is technological progress in the short-run.

JEL classification codes: O30, O47, O57, C22.
Key words: economic growth, technological progress, the bounds testing approach,

ARDL, FM-OLS

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I. Introduction

This study explores sources of economic growth for Hong Kong, Republic of Korea, Singapore and Taiwan, contributes to the debate over whether the sources of economic growth stem from technological progress or capital accumulation in East Asian economies, and deliberates on the identifying assumption generally used in growth accounting studies. Traditional sources of growth studies assume the nature of technological progress to be Hicks-neutral (Solow 1957: 312; Barger 1969: 144; Nishimizu and Hulten 1978: 352; de Gregorio 1992: 64; Senhadji 2000: 132; Altug, Filiztekin and Pamuk 2008: 403; Fuentes, Larrain and Schmidt-Hebbel 2006: 121; Abu-Qarn and Abu-Bader 2007: 753; van der Eng 2010: 295). The present study argues against this assumption. Although several studies based on time series econometrics implicitly or explicitly assume that there are long-run equilibrium relationships and steady state conditions, they also assume Hicks-neutral technological progress (Senhadji 2000; Abu-Qarn and Abu-Bader 2007). However, if there are steady state conditions, the nature of technological progress should be assumed to be Harrod-neutral (see Uzawa 1961).

This study considers the economies of Hong Kong, the Republic of Korea, Singapore and Taiwan, known as the "East Asian Tigers". These economies enjoyed a remarkable record of high and sustained economic growth over three decades from the mid-1960s to the early 1990s. Their ability to achieve fast economic growth has led many economists to wonder whether it stems from capital accumulation or technological progress. Collins and Bosworth (1996) emphasise that East Asian economies are distinguished by the magnitude of their capital accumulation and that the contribution of productivity is quite ordinary. Young (1992, 1994, 1995) and Kim and Lau (1994) suggest that productivity growth in East Asia is unimportant and that the main source of growth is capital accumulation. Park and Ryu (2006) show that physical capital accumulation is an important source of economic growth in East Asian economies when a homothetic function is used, whereas it is technical progress in the Cobb-Douglas production function with constant returns to scale. Klenow and Rodríguez-Clare (1997) report that technological progress account for the most growth in output per worker in Hong Kong, the Republic of Korea, and Taiwan.

This study asks the following question: What are the theoretical and empirical results of assuming the nature of technological progress as Harrod-neutral in growth accounting for the four "East Asian Tigers"?

From this perspective, the elasticity of the output per labour for the capital stock per labour are first estimated for Hong Kong, the Republic of Korea, Singapore and Taiwan, for the period between the 1950s and 2007. Second, the contributions of technological progress and capital accumulation to per labour output growth are decomposed, using the short-term coefficients obtained from long-term information.

The paper is organised as follows. Section II discusses the theoretical background of the paper. Section III gives the methodology of the study. Section IV gives information on the data set. Section V includes the model estimation results and discusses the findings. Section VI concludes the paper.

II. Theoretical background

As Uzawa (1961) proved, the nature of technological progress consistent with steady state conditions is Harrod-neutral (see also Allen 1967). Here, steady state indicates a long-run equilibrium relationship. If long-run equilibrium relationships are analyzed, the nature of technological progress should be assumed to be Harrod-neutral rather than Hicks-neutral, i.e., the production function should be assumed as $Y = F(K, AL) = K^{\alpha}(AL)^{1-\alpha}$ rather than $Y = F(AK, AL) = AK^{\alpha}L^{1-\alpha}$, where Y, L and K indicate the level of output, labour and capital stock, respectively, and A indicates the level of technology. Furthermore, the time series econometrics analysis is generally based on testing whether there is a long-run equilibrium relationship among the non-stationary variables. If so, sources of growth studies should assume that the nature of technological progress is Harrod-neutral. Gundlach (2005) addresses this problem and reports that traditional growth accounting studies usually assume the nature of technological progress to be Hicks-neutral rather than Harrod-neutral, although the Harrod-neutral technological progress is the most convenient identifying assumption to use in an empirical long-run growth analysis. As Gundlach emphasises (2005: 544), the identifying assumption is the one that “determines how the underlying production function might shift”. If the nature of the technological progress is Harrod-neutral, the capital-output ratio is kept constant, and steady state conditions are allowed.

Conversely, as Romer (1996) emphasises, the sources of growth method is used to analyse short-run growth. As Romer (1996: 26) states briefly, “In the Solow model, long-run growth of output per labour depends only on technological progress. But short-run growth can result from either technological progress or capital accumulation. Thus the model implies that determining the sources of short-run growth is an empirical issue.” According to our discussion, this empirical issue should depend on a long-run analysis if time series econometrics methods are used. A contradiction thus arises between economic and econometric analysis.

The econometric dimension of this contradiction is first briefly explained. Economic variables pertain to a system. After a shock, there may be disequilibrium among economic variables in the system in the short-run. This disequilibrium may be compensated, and the system may return to the equilibrium state in the long-run. To analyse the system correctly and test the relationships among variables predicted by the theory, information on the long-run should be preserved.

If the system moves away from the equilibrium, economic variables move away from the attractor. In this paper, theoretically speaking, the capital-output ratio can be considered the attractor. According to the neoclassical growth theory, any shock leads to a movement away from the capital-output ratio, and the system is programmed to move towards the attractor. However, how can the existence of such an attractor be econometrically proved? How can one impose the tendency of moving away from or towards the attractor due to a shock onto an econometric model? Engle and Granger (1991) define two types of economic series, long-memory (non-stationary) and short-memory (stationary), to answer these questions. In a short-memory system, an old shock to the series has virtually no effect on the current value of the series if the shock happened long enough ago. In a long-memory system, an old shock has a noticeable impact on the current value of the series. A long-memory series has no attractor, and there is generally no tendency to return to any value, including its initial value. Suppose that we have two long-memory series (x_t and y_t) and changes in y_t are explained by changes in x_t. This causes the well-known spurious regression problem. In this case, two series are spuriously related because they are both trended and estimators, and test statistics may be misleading (Granger and Newbold 1974). The series, which includes linear combinations of two or more such series, may be a short-memory series. In such a case, estimators and test statistics may continue to preserve their validities. This situation is called cointegration and is a sufficient condition for the existence of an attractor; it can also correspond to the equilibrium that arises in macroeconomic theory.

Granger (1993: 307) summarises the implications of these explanations thus: “if macro theories are about equilibria, econometric techniques are not, it becomes difficult for these theorems to be tested on actual data”. In a time series analysis, researchers first test whether their series are long-memory or stationary series. Because most economic time series are non-stationary, multivariate time series analysis focuses on the cointegration relationships among them.

As Pesaran (1997: 178) mentions, "the notion of long-run is inextricably linked with the concept of equilibrium in economics, although in much time series econometrics long-run analysis is conducted without providing an explicit account of the type of equilibrium theory that may underlie it." Time series analysis does not clearly consider information about the type of equilibrium predicted by economic theory. Cointegration analysis inspires this non-theoretical approach for modelling the long-run. Thus, we must draw attention that there is an important distinction between the model used to estimate long-run levels relationship and that introduced by growth accounting.

Equation (1) gives the estimation equation usually used in sources of growth studies.

$$\ln(Y/L)_t = C + \alpha \ln(K/L)_t + u_t \ , \tag{1}$$

where C and u are the constant and error terms, respectively, Y is the level of output, L is the labour and K is the capital stock. The attractor is the error-term for the series of $\ln(Y/L)_t$ and $\ln(K/L)_t$. If the nature of the technological progress is assumed to be Harrod-neutral, the constant term equals $C = (1-\alpha)\ln A$, while it is $C = \ln A$ if the nature of the technological progress is Hicks-neutral.[1]

If equation (1) is rearranged according to the growth rates assuming Harrod-neutral technological progress and the error-term is ignored, it is expressed by equation (2):

$$\frac{dA}{dt}\frac{1}{A_t} = \frac{d(K/L)_t}{dt}\frac{1}{(K/L)_t} = \frac{d(Y/L)_t}{dt}\frac{1}{(Y/L)_t} \ . \tag{2}$$

[1] By definition, Hicks-neutral technological progress occurs if the capital-labor ratio does not change while the ratio of factor prices is constant (Hicks 1963: 121). Conversely, Harrod-neutral technological progress occurs if the capital-output ratio does not change while the marginal productivity of per labor capital stock is constant (Harrod 1948: 82). If Hicks-neutral technological progress occurs, income grows at a certain rate in the short-run. However, if Harrod-neutral technological progress occurs, income grows at a rate equals to the growth rate of technological progress in the long-run.

In the long-run, since the growth rate of capital per labour is equal to the growth rate of technological progress, the parenthesis in equation (2) is equal to zero. Thus, the growth rate of output per labour is equal to the growth rate of technological progress, and this justifies the steady state conditions; i.e. the growth rate of output per labour and capital per labour are equal to the growth rate of technological progress:

$$\frac{dA}{dt}\frac{1}{A_t}=\frac{d(K/L)_t}{dt}\frac{1}{(K/L)_t}=\frac{d(Y/L)_t}{dt}\frac{1}{(Y/L)_t}\ . \tag{3}$$

Equation (3) is compatible with the neoclassical growth model and long-run economic growth. If it is assumed that the nature of the technological progress is Hicks-neutral, then the constant term would equal $C=\ln A$. The following equation can thus be written:

$$\frac{d(Y/L)_t}{dt}\frac{1}{(Y/L)_t}=\alpha\frac{d(K/L)_t}{dt}\frac{1}{(K/L)_t}+\frac{dA}{dt}\frac{1}{A_t}\ . \tag{4}$$

As the growth rate of output and capital per labour equal the growth rate of technological progress in the long-run, equation (4) is not compatible with long-run growth conditions.

Assuming that there are constant returns to scale conditions and the nature of the technological progress is Harrod-neutral, the production function can be written in the Cobb-Douglas form:

$$Y_t=K_t^{\alpha}\left(A_tL_t\right)^{1-\alpha}\ . \tag{5}$$

Note that it is assumed that the level of technology includes the human capital stock, and the measure of the human capital stock is embodied in the labour force, so the expression (A_tL_t) denotes a skill-adjusted measure of the labour input. Since the human capital stock is a comprehensive concept, i.e., it includes investment in health, education, etc., it is reasonably possible to assume the human capital as an input being embodied in labour rather than assuming its components as separate inputs. Moreover, it is also assumed that not only human capital stock but also all of the other components that have an impact on the level of technology or productivity are embodied in the labour force. However, this assumption suggests that the empirical results should be interpreted carefully. If equation (5) is rearranged, equation (6) can be written as follows:

$$\left(\frac{Y}{L}\right)_t = A_t\left(\frac{K}{Y}\right)_t^{\frac{\alpha}{1-\alpha}}. \tag{6}$$

If the natural logarithm of equation (6) is taken, the estimation equation can be written as follows:

$$\ln(Y/L)_t = C + \beta\ln(K/Y)_t + u_t \tag{7}$$

where C, $C = \ln A_t$, and u are the constant and error terms, respectively, Y is the level of output, L is the labour and K is the capital stock, and $\beta = \alpha/(1-\alpha)$. If equation (7) is rearranged using the growth rates and the error-term is ignored, it is expressed by equation (8), which gives the final equation for growth accounting:

$$\frac{d(Y/L)_t}{dt}\frac{1}{(Y/L)_t} = \frac{dA}{dt}\frac{1}{A_t} + \beta\frac{d(K/Y)_t}{dt}\frac{1}{(K/Y)_t}. \tag{8}$$

We must stress that short-run parameters should be estimated to determine the sources of short-run growth, i.e., the contribution of capital accumulation and technological progress. Thus, we first estimate the long-term coefficients for the relevant countries. We then estimate the short-term coefficients depending on the long-term estimation results. Based on the short-term coefficients, we calculate the contribution of the technological progress and capital stock.

III. Econometric methodology

In the empirical analysis, the autoregressive distributed lag (ARDL) approach of Pesaran and Shin (1999) is employed to estimate the long-run relationships between gross domestic product per labour and the capital-output ratio. The bounds testing approach to cointegration, developed by Pesaran, Shin, and Smith (2001), is used to test the existence of a long-run relationship. Compared to the other tests, i.e., the two-stage estimation of Engle and Granger (1987) and the full information method of Johansen (1988) and Johansen and Juselius (1990), the bounds testing approach can be applied irrespective of whether the underlying regressors are purely $I(0)$, purely $I(1)$, fractionally integrated, or mutually co-integrated.

The ARDL approach is a two-stage approach and, as in Pesaran and Shin (1999), such a two-stage procedure has two important advantages: (i) it effectively corrects for possible endogeneity of explanatory variables, and (ii) the estimates exhibit desirable small sample properties (see also Panopoulou and Pittis 2004; Caporale and Pittis 2004).

Equation (7) gives the empirical model specification that relates the real gross domestic product per labour ($\ln(Y/L)$) and the ratio of physical capital stock to the gross domestic product ($\ln(K/Y)$). Without having any prior information about the direction of the long-run relationship among the variables, the bounds testing approach estimates an unrestricted error-correction model (UECM), taking each variable in turn as a dependent variable. Equation (9) gives a general form for this unrestricted model of $\ln(Y/L)$ on $\ln(K/Y)$.

$$\Delta\ln(Y/L)_t = c_0 + \delta_1 \ln(Y/L)_{t-1} + \delta_2 \ln(K/Y)_{t-1} + \sum_{j=1}^{p} \lambda_j \Delta\ln(Y/L)_{t-j} + \sum_{j=1}^{p} \omega_j \Delta\ln(K/Y)_{t-j} + \psi D_t + \delta t + u_t , \quad (9)$$

In equation (9), D_t is a vector of exogenous variables, e.g., structural change dummies, and Δ indicates a first difference operator. The first stage in the bounds testing approach is to estimate equation (9) using ordinary least squares (OLS). According to the model, the null hypothesis of no co-integration ($\delta_1 = \delta_2 = 0$) against the alternative of a long-run levels relationship ($\delta_1 \neq \delta_2 \neq 0$) is performed as a Wald restriction test. "The test statistic underlying the procedure is the familiar Wald or F-statistic in a generalised Dickey-Fuller type regression used to test the significance of lagged levels of the variables under consideration in a conditional unrestricted equilibrium correction model" (Pesaran, Shin and Smith, 2001, p.

289-290). The asymptotic distributions of the *F*-statistics are non-standard under the null hypothesis of no co-integration between the variables in the UECM given in equation (9), whether the variables are purely $I(0)$, purely $I(1)$, or mutually co-integrated.

Pesaran, Shin and Smith (2001: 300-301, 303-304) provide two sets of asymptotic critical values. In the first and second sets, all variables are assumed to be $I(0)$ and $I(1)$, respectively. The null hypothesis of no co-integration can be rejected, indicating that a long-run equilibrium exists among the variables if the computed *F*-statistics is greater than the upper bound critical value. The null hypothesis of no co-integration can be accepted if the computed *F*-statistics is less than the lower bound of the critical value, and the bounds test is inconclusive if the computed *F*-statistics falls within the lower (first critical value set) and upper bounds (second critical value set) of the critical values.

A two-step procedure is followed if a long-run relationship is established in the first stage. In the second stage, a conditional autoregressive distributed lag model, ARDL(p,q), for $\ln(Y/L)$ can be estimated as follows:

$$\ln(Y/L)_t = c_0 + \sum_{j=1}^{p} \gamma_j \ln(Y/L)_{t-j} + \sum_{j=0}^{q} \theta_j \ln(K/Y)_{t-j} + \psi' D_t + u_t . \tag{10}$$

Here, all variables are defined as above, and the lag lengths p and q relating to the two variables in the model are selected using the Akaike (AIC) or Schwarz Bayesian (SIC) Information Criteria. The long-run parameters $\hat{C}$ and $\hat{\beta}$ in equation (7) can easily be obtained from the OLS estimates of equation (10), thus:

$$\hat{C} = \frac{\hat{c}_0}{1 - \sum_{k=1}^{p} \hat{\gamma}_k} \text{ and } \hat{\beta} = \frac{\sum_{j=0}^{q} \hat{\theta}_j}{1 - \sum_{k=1}^{p} \hat{\gamma}_k} . \tag{11}$$

The second step in the second stage of the bounds testing ARDL approach involves estimating a conditional ECM model. “A principle feature of cointegrated variables is that their time paths are influenced by the extent of any deviation from long-run equilibrium. After all, if the system is to return to long-run equilibrium, the movements of at least some of the variables must respond to the magnitude of disequilibrium” (Enders, 2004: 328). Equation (12) specifies the conditional ECM model.

$$\Delta \ln(Y/L)_t = \mu + \sum_{j=1}^{p} \lambda_j \Delta \ln(Y/L)_{t-j} + \sum_{j=0}^{q} \omega_j \Delta \ln(K/Y)_{t-j} + \vartheta ECM_{t-1} + \psi' D_t + \delta t + u_t , \qquad (12)$$

where λ_j and ω_j are short-run parameters, ϑ is the speed of adjustment, which determines model’s convergence to equilibrium, and the error-correction term ECM_t is defined as

$$ECM_t = \ln(Y/L)_t - \hat{C} - \hat{\beta} \ln(K/Y)_t . \qquad (13)$$

While the ARDL approach is used for the main results of the paper, a second estimator is also employed to test the sensitivity of the results in the estimator choice. The second estimator used in this paper is the fully modified ordinary least square (FM-OLS) estimator of Phillips and Hansen (1990). The FM-OLS estimator has two direct advantages: (i) it corrects for endogeneity and serial correlation effect, and (ii) it asymptotically eliminates the sample bias.

The estimated parameters may change over time due to structural change(s) in the data generating processes. This paper applies parameter stability tests as cointegrated regression models proposed by Hansen (1992a). Hansen (1992a) extends the FM-OLS estimator to cover general models with stochastic and deterministic trends. The author proposes three statistics, i.e., Sup*F*, Mean*F*, and L_C, to test the null hypothesis that the estimated long-run parameters are stable. However, their alternative hypotheses are different. The Sup*F* test dates back to Quandt (1960), and its alternative hypothesis is a swift shift in regime. The Mean*F* test is appropriate when the question under investigation is whether the specified model captures a stable relationship. The Sup*F* and Mean*F* tests require truncating the sample size. We use the subset [0.10*T*, 0.90*T*] because we have small samples; *T* denotes sample size. The L_C statistic is recommended if the parameter variation likelihood is relatively constant throughout the sample. As Hansen (1992a) notes, the L_C test can be used to test the null cointegration against the alternative of no cointegration.

Some instability problems could result from insufficient modelling of the short-run dynamics that describe departures from the long-run equilibrium

(Bahmani-Oskooee 2001). We have thus tested stability in the ECMs by applying the testing procedure of Hansen (1992b). He describes a simple and powerful test for parameter instabilities in dynamic models like ECMs. We can use this either to test the stability of each parameter individually or to test the stability of all parameters jointly. Moreover, Hansen (1992b) presents a test statistic for testing the stability of the error-variance.

IV. Data set

The data set covers the period 1951-2007 for Taiwan, 1953-2007 for the Republic of Korea and 1960-2007 for Hong Kong and Singapore. Data have been obtained from the Penn World Table Version 6.3 (PWT 6.3). The dependent variable of the model is the real gross domestic product per labour (Y/L). We must perform several transformations to obtain these series because they were not directly obtained from Penn World TableVersion 6.3. First, the total real GDP is obtained from the real GDP per capita (*rgdpl2*) by multiplying it by the population (*pop*). Second, the number of workers is obtained from the real GDP per worker series (*rgdpl2wok*).[2] The real GDP per labour series is obtained by dividing the total real GDP by the number of workers. Hereafter, we use the real GDP per worker term instead of the real GDP per labour.

The independent variable of the model is the capital-output ratio (K/Y). Though the ARDL approach of Pesaran and Shin (1999) and the FM-OLS approach of Philips and Hansen (1990) effectively correct for possible endogeneity in the explanatory variables, we use the investment share of real GDP (I/Y) instead of (K/Y) to avoid possible endogeneity in the explanatory variables.

One explanation for using (I/Y) instead of (K/Y) is to assume (K/Y) is constant over the examination period. However, if we initially assume (K/Y) is constant, we cannot calculate the contribution of capital accumulation which occurs because of a change in (K/Y). Our explanation for using (I/Y) instead of (K/Y) is as follows.

[2] As noted in Penn World Table Version 6.3, workers include all status categories of persons in employment, not only employees (including paid family workers) but also employers, own-account workers, members of producer cooperatives, contributing family workers and workers not classifiable by status (Heston, Summers and Aten 2009).

As shown by Mankiw, Romer and Weil (1992) and Klenow and Rodríguez-Clare (1997), the capital-output ratio equals

$$\frac{K}{Y}=\frac{I/Y}{n+g+\delta} \tag{14}$$

where I is investment, n is the growth rate of labour, g is the growth rate of technology and δ is the depreciation rate. However, this equation is valid only in steady state conditions where $\frac{Y}{AL}=\left(\frac{s}{n+g+\delta}\right)^{\frac{\alpha}{1-\alpha}}$ and $\frac{K}{AL}=\left(\frac{s}{n+g+\delta}\right)^{\frac{1}{1-\alpha}}$. Indeed, if the steady state value of (K/AL) is divided by the steady state value of (Y/AL), $\frac{K}{Y}=\frac{s}{n+g+\delta}$ is obtained, and since $s=\frac{I}{Y}$, then $\frac{K}{Y}=\frac{I/Y}{n+g+\delta}$ is valid in *steady state*. Thus, if (K/Y) is proxied by $\frac{I/Y}{n+g+\delta}$, it is assumed that every observed value of (K/Y) is equal to the steady state value of it. However, when the capital-output ratio changes, the economy is out of the steady state. To us, it is reasonably more realistic and practical to assume that every observed value of (K/Y) is different from its steady state value in proportion to its steady state parameters. In other words, it is assumed that $\frac{K}{Y}=\frac{I/Y}{n+g+\delta}\gamma$ and $\gamma=n+g+\delta$. Simply we assume that $\gamma=n+g+\delta$ because this assumption guarantees that every economy would be assessed according to its own steady state parameters. As an example, we suppose that there are two economies and each of them has its own steady state parameters:

$$\left(\frac{K}{Y}\right)_1=\frac{(I/Y)_1}{n_1+g_1+\delta_1} \tag{15}$$

$$\left(\frac{K}{Y}\right)_2=\frac{(I/Y)_2}{n_2+g_2+\delta_2}. \tag{16}$$

For each observed period, the first and the second economy has a (K/Y) value which is different from their own steady state (K/Y) value, in proportion to

$n_1 + g_1 + \delta_1$ and $n_2 + g_2 + \delta_2$, respectively:

$$\left(\frac{K}{Y}\right)_1 = \frac{(I/Y)_1}{n_1 + g_1 + \delta_1}\gamma_1 , \tag{17}$$

$$\left(\frac{K}{Y}\right)_2 = \frac{(I/Y)_2}{n_2 + g_2 + \delta_2}\gamma_2 , \tag{18}$$

where $\gamma_1 = n_1 + g_1 + \delta_1$ and $\gamma_2 = n_2 + g_2 + \delta_2$.

Moreover it is assumed that the difference between the steady state value and the observed value of (K/Y) is determined by each economy's own steady state parameters. Obviously, this assumption makes it possible and simple that the observed (I/Y) values can be used as a proxy for (K/Y).

Penn World Table Version 6.3 gives the investment share of real GDP per capita (as a percentage, and its code is *ki*). The (I/Y) series is obtained by manipulating the investment share of a real GDP per capita series by population. Finally, all data except the number of workers are in 2005 US dollars and logs.

V. Empirical results

As Pesaran, Shin and Smith (2001) indicate, the ARDL approach to test the existence of a relationship between variables in levels is irrespective of whether the underlying regressors are purely $I(0)$, purely $I(1)$, or mutually cointegrated. Pre-testing can lead to bias in estimated cointegrating vector or vectors, and it creates uncertainties regarding the cointegration inference, as it depends on the significance level chosen with unit root tests (Maddala and Kim 1998). However, we determine the order of integration of the series by employing several unit root tests with and without structural breaks.[3] Given that one series might be $I(1)$ and

[3] To determine the data integration order, the augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1979), the generalised least squares detrended Dickey-Fuller (DF-GLS) test (Elliot, Rothenberg, and Stock 1996) and the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test (Kwiatkowski, Phillips, Schmidt, and Shin 1992) were used. To check the effect of structural changes on the unit root test results, we also used the tests developed by Zivot and Adrews (1992) and Lee and Strazicich (2003). These test results show that all series in our dataset are either I(1) or I(0), but not I(2). Given that some series might be I(1) and others I(0), the mixed evidence on the order of the series, we proceed to test for long-run levels relationship using the bounds tests.

the other $I(0)$, the mixed evidence on the order of integration of the series, we proceed to test for a long-run levels relationship using the bounds test.

Before the empirical results, points in the bounds testing analysis will be discussed. As can be seen from Figures 1-3, the real GDP per worker series of each country steadily rises over time. For example, Figure 1 shows the real GDP per worker series of Hong Kong increases at a decreasing rate. These suggest, at least initially, that the real GDP per worker equation should include a linear trend. We estimate the test regression given in equation (9) by OLS with and without a linear time trend to determine the appropriate lag length p and whether a deterministic linear trend is required in addition to the $\ln(Y/L)$ series, for $p = 1, 2, \ldots, 6$.

The specification given in equation (9) is based on the assumption that the disturbances u_t are serially uncorrelated. Therefore, it is important that the lag order p of the test regression is selected appropriately (see Pesaran, Shin and Smith 2001). The appropriate lag for the test regression with or without linear deterministic trend are selected both using AIC and SIC, by controlling residual serial correlation against first and fourth order. We also checked whether the lagged changes of the real GDP per worker and the lagged changes of the investment share of real GDP are significant in the test regressions while determining the appropriate lag length. Thus, we avoid unnecessary over-parameterisation.

We calculate three F-statistics (*i*) under the constraint of unrestricted intercept, and with no trend (F-iii); (*ii*) under the constraint of unrestricted intercept, and with restricted trend (F-iv) and under the constraint of unrestricted intercept, and with unrestricted trend (F-v).[4] According to the test results, there is a long-run levels relationship between the $\ln(Y/L)$ and $\ln(I/Y)$ series at least at the 5% significance level for all the cases. Details on unit root tests results, lag length selection and bounds test results are available on request.

Empirical results for each country are separately given in this section, beginning with a brief economic history.[5]

[4] Narayan (2004, 2005) examines the small sample problem within the context of the bounds testing approach. He generates the critical values for the test statistics (the F-statistics) to accommodate small sample sizes. The present paper uses the critical values of the F-statistics modified by Narayan (2005), as we have 48-57 observations (T).

[5] We also estimated sources of growth model for Singapore which is identified among eight high-performing Asian economies in World Bank (1993). However, all of the estimated values of α in Equation (6) for Singapore were not compatible with the assumption that the value of α is between zero and one. Therefore, the growth accounting results were not reasonable. Both long and short-run estimates for Singapore are not reported here, but they are available on request.

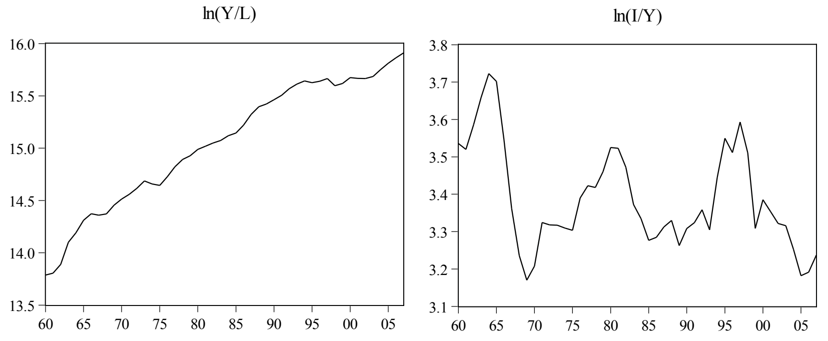
A. Empirical results for Hong Kong

Hong Kong, as a British colony, became the Asian headquarters of British and firms of other nations in the 1840s. Hong Kong has been the pivotal meeting place in Asia and in foreign social capital networks since the late nineteenth century. After World War II, the Chinese Communist Party provided an entrepreneurial base for the city to grow as a major export industrial centre. Trading and financial firms in Hong Kong gave manufacturers access to world markets (Meyer 2008). After reforms in 1978, industrialists in Hong Kong began their move into Guangdong province. The government in Hong Kong has followed relatively little industrial policy (Jomo, 2001), and its economy has been transformed from an export-oriented manufacturing economy to a service-dominated economy (Meyer 2000 and Chiu, Ho, and Lui 1997).

The modernisation reforms of the People's Republic of China (PRC) have allowed the Hong Kong manufacturing industry to migrate to the Pearl River Delta in mainland China, where low land and labour costs are well suited to manufacturing. This has transformed the Hong Kong economy into one dominated by trade-related services (Tao and Wong 2002). The continued inflow of foreign direct investment in the finance and producer service sectors has helped cement the status of Hong Kong as a global city. Developing stock and finance markets in Hong Kong generated a large inflow of foreign portfolio investments, which in turn fostered its development as the financial centre of the region in the 1990s (Fung and Hung 2010).

Figure 1 below shows real GDP per worker and the investment share of the real GDP of Hong Kong over a period of forty-eight years. While the investment share of real GDP displays upward and downward movements, the real GDP per worker series increases at a decreasing rate. The investment share of real GDP reached its highest values in 1965 and 1997 according to Penn World Table Version 6.3. As a financial centre in Asia, Hong Kong was greatly affected by the Asian financial crisis in 1998, and the economy contracted by 5.6%.

Figure 1. Hong Kong: real GDP per worker and investment share of real GDP



Note: the arguments of the log function are multiplied by 100.

The Sino-British Accord in 1984, agreeing to return Hong Kong to China in 1997, contributed to a politically sensitive economic environment and an uncertain political climate. To account for both the effect of this political and economical change and the Asian financial crisis, the D_t vector is specified, containing one dummy variable: $D97 = 1$ for the period 1997-2007, 0 otherwise.

To obtain the long-run coefficient estimates of the real GDP per worker equation, we first estimate the ARDL(p,q) model using OLS. The appropriate lag length is chosen using both AIC and SBC. Long-run estimates are obtained from the selected ARDL models given in Equations (10) and (11). ARDL models are also estimated using linear deterministic trends. The standard errors of the estimated long-run parameters are computed using the delta-method. [6]

As Table 1A shows, significant estimates are obtained from ARDL models, including the linear trend term. ARDL estimates of ln$A$ and $\alpha/(1-\alpha)$ are positive and statistically significant at the conventional significance levels. Using these estimates, the respective elasticity coefficients (α) are estimated to be 0.19 and 0.23. The FM-OLS estimation results for the long-run relationship show that the response of $\ln(I/Y)$ to $\ln(Y/L)$ is negative and significant at the 1% level. FM-OLS produces different results and a different elasticity coefficient (approximately 1.1). While FM-OLS estimates the elasticity coefficient greater than 0.5, ARDL estimates this coefficient to be less then 0.5.

[6] The delta method is used to estimate the standard error of a parameter, which is the non-linear combination of the other parameters (Green, 2008).

Table 1. Hong Kong

A. Long-run estimates of the real GDP per worker equation with ARDL and FM-OLS

Dependent variable: ln(Y/L)	Without deterministic trend			With deterministic trend		
	AIC[a]	SBC[b]	FM-OLS	AIC[a]	SBC[b]	FM-OLS
Constant (C)	18.449 (3.805)***	18.449 (3.805)***	49.559 (7.947)***	13.354 (29.755)***	13.132 (31.249)***	50.344 (9.381)***
Coefficient on ln(I/Y) [β]	-0.666 (-0.450)	-0.666 (-0.450)	-10.140 (-5.509)***	0.238 (1.840)*	0.299 (2.433)**	-10.417 (-6.577)***
Estimate of α in eq. (6)	-1.994	-1.994	1.109	0.193	0.230	1.106
Stability tests of Hansen (1992a)						
Sup$F$		18.605			23.662	
Mean$F$		4.380♣			6.113♣	
L_c		0.560♣			0.815♣	

Note: the SupF and MeanF statistics are calculated using the trimming region [0.10, 0.90]. For stability tests, the null hypothesis is that "the estimated long-run parameters are stable".The Bartlett kernel is used for bandwidth to estimate the elements of covariance matrix. See Hansen (1992a: 327-328) for the critical values. t-statistics are given in parentheses. The symbols ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively. ♣ shows that the null hypothesis cannot be rejected at the 1% level. [a] The selected ARDL model is (2,1) and (3,1) by AIC. [b] The selected ARDL model is (2,1) and (3,0) by SBC.

B. Short-run Estimates for the real GDP per worker equation

Dependent variable: Δln(Y/L)	AIC		SBC	
	Without deterministic trend	With deterministic trend	Without deterministic trend	With deterministic trend
Intercept	0.005 (0.389)	0.014 (1.694)*	0.005 (0.389)	0.015 (1.945)*
Δln$(Y/L)_{t-1}$	0.207 (1.573)	0.137 (1.165)	0.207 (1.573)	0.132 (1.105)
Δln$(Y/L)_{t-2}$	-	-0.125 (-1.142)	-	-0.137 (-1.252)
Δln$(I/Y)_t$	0.181 (2.357)**	0.168 (2.613)**	0.181 (2.357)**	0.174 (2.699)***
$D97$	-0.063 (-1.602)	-0.078 (-2.525)**	-0.063 (-1.602)	-0.085 (-2.731)***
ECM_{t-1}	-0.029 (-2.592)***	-0.340 (-5.871)***	-0.029 (-2.592)***	-0.339 (-5.822)***
R^2	0.375	0.619	0.375	0.615
F-stat.	6.157***	12.646***	6.157***	12.489***
DW	1.918	1.789	1.918	1.780
Hansen stability tests for the ECMs				
Joint L_c	0.865 [0.590]	0.871 [0.760]	0.865 [0.590]	0.810 [0.820]
Error-variance	0.258 [0.170]	0.099 [0.570]	0.258 [0.170]	0.093 [0.600]

Note: t-statistics are given in parentheses. p-values are given in brackets. The symbols ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively. For stability tests, the null hypothesis is that "the estimated short-run parameters are stable". See Hansen (1992b) for the critical values.

According to the parameter stability test results reported in Table 1A, there is strong evidence that long-run parameters are relatively constant through time, as the calculated values for the Mean*F* and L_C tests are smaller than their critical values at the 1% level. The results for the L_C test also confirm that the real GDP per worker maintains a long-run relationship with the real investment share of real GDP.

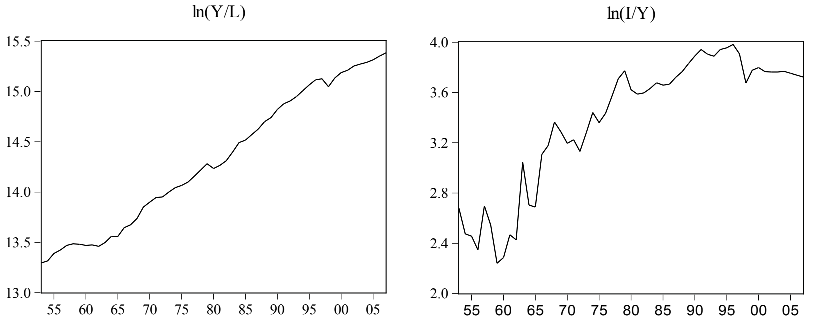
The first and third columns in Table 1B give the short-run parameter estimates of the error-correction models obtained from ARDL(2,1) without linear deterministic trends. In this equation, the coefficient of $\Delta \ln(Y/L)_{t-1}$ is insignificant. When its parameter is restricted by zero, the calculated Wald *F*-statistic is 2.475 with a *p*-value of 0.12. The ARDL(3,1) model selected by the AIC and the ARDL(3,0) model selected by SBC include $\Delta \ln(Y/L)_{t-1}$ and $\Delta \ln(Y/L)_{t-2}$. In both specifications, we also restricted the parameters to zero. The calculated *F*-statistics are 0.953 with a 0.40 *p*-value and 1.009 with a 0.37 *p*-value. When we restrict the models, the $\Delta \ln(Y/L)_t$ coefficient is estimated to be significant between 0.195 (at the 5%) and 0.209 (at the 1%), and the coefficient of ECM term is estimated to be significant between −0.028 and −0.213. Coefficient of the dummy variable is also estimated significantly between −0.065 and −0.085. To be consistent with the models in Table 1A, we report the estimates obtained from the selected ARDL models. In the last part of this section (in Table 4), we use the estimates given in this paragraph.

Following Hansen (1992b), here we report the stability test results to check the stability of the error-variance and that of all parameters jointly. The joint stability test (L_C) results indicate that the null hypothesis of the parameter stability cannot be rejected at the 1% significance level in all specifications (Table 1B). Similar results are obtained for the error-variance at the 1% significance level. These stability test results are also obtained when we restrict the insignificant coefficients to zero.

B. Empirical results for the Republic of Korea

As Figure 2 shows, the two series display an increasing trend over time. There are three noteworthy structural changes in the Korean economy. Following the Korean War (1950-53), the economy experienced a slow economic recovery between 1953 and 1961. The period from 1962-1971 is characterised by economic reforms emphasising labour-intensive light manufacturing exporting industries. From 1972-1979, the strategy of promoting heavy and chemical industries was followed, and it resulted in rapid monetary expansion and increased budget deficits. The 1980s and 1990s experienced major shifts in economic policy, including economic stabilisation and liberalisation and an opened economy. The 1990s witnessed the onset of a financial crisis (1990-1997) because of the integration into the global economy. However, economic growth remained strong during this period.

Figure 2. Republic of Korea: real GDP per worker and investment share of real GDP



Note: the arguments of the log function are multiplied by 100.

After the 1997 financial crisis and the economic collapse of 1998, Korea instituted reforms and reconstructed the financial and corporate sectors. The economic growth rate was recorded as approximately 6% between 2000 and 2004 (Harvie and Pahlavani, 2006; OECD, 2005; Harvie and Lee, 2003a, 2003b and Smith, 2000). To consider the effect of these structural changes, the D_t vector is specified containing three dummy variables: D5361 = 1 for the period 1953-1961, 0 elsewhere; D7279 = 1 for the period 1972-1979, 0 elsewhere; and D9097 = 1 for the period 1990-1997, 0 elsewhere.

Table 2A reports the estimates of the long-run levels equation with the parameters obtained using the ARDL and FM-OLS approaches. The ARDL estimates for $\ln A$ and $\alpha/(1-\alpha)$ are positive and statistically significant. These results are also obtained from the selected ARDL(1,0) model, including trend terms. The elasticity coefficient is estimated to be 0.63 by manipulating the $\alpha/(1-\alpha)$ estimates. However, the estimates for the $\alpha/(1-\alpha)$ parameter obtained from the ARDL with trend models are smaller, and $\hat{\alpha}=0.21$.The FM-OLS estimates of the real GDP per worker equation are similar in their estimated parameter signs; however, the $\alpha/(1-\alpha)$ coefficient estimates are smaller for the Republic of Korea. Using the FM-OLS $\alpha/(1-\alpha)$ coefficient estimates produces a similar result as using the ARDL model with constant ($\hat{\alpha}\cong 0.60$).

Table 2. Republic of Korea

A. Long-run Estimates of the real GDP per worker equation with ARDL and FM-OLS

Dependent variable: ln(Y/L)	Without deterministic trend			With deterministic trend		
	AIC[a]	SBC[b]	FM-OLS	AIC[a]	SBC[b]	FM-OLS
Constant	4.816 (4.455)*	4.809 (5.225)*	5.999 (19.358)***	7.846 (32.328)*	7.846 (32.328)*	6.235 (19.557)***
Coefficient on ln(I/Y)	1.685 (5.204)*	1.654 (6.071)*	1.511 (26.850)***	0.267 (2.767)*	0.267 (2.767)*	1.476 (25.499)***
Estimate of α in eq. (6)	0.627	0.623	0.601	0.210	0.210	0.596
Stability tests of Hansen (1992a)						
Sup$F$		325.879			181.528	
Mean$F$		32.096			19.288	
L_c		0.293♣			0.761♣	

Note: see Table 1A. [a] The selected ARDL model is (1,1) and (1,0) by AIC. [b] The selected ARDL model is (1,0) and (1,0) by SBC.

B. Short-run estimates of the growth rate of the real GDP per worker equation

Dependent variable: $\Delta\ln(Y/L)$	AIC		SBC	
	Without deterministic trend	With deterministic trend	Without deterministic trend	With deterministic trend
Intercept	0.001 (0.094)	0.015 (2.306)**	0.005 (0.587)	0.015 (2.306)**
$\Delta\ln(I/Y)_t$	0.101 (3.963)***	0.085 (3.693)***	0.099 (3.937)***	0.085 (3.693)***
$D5361$	0.034 (1.326)	0.070 (2.760)***	0.034 (1.342)	0.070 (2.760)***
$D7279$	0.028 (1.585)	0.035 (2.003)**	0.028 (1.615)	0.035 (2.003)**
D9097	0.058 (3.159)***	0.068 (3.861)***	0.058 (3.119)***	0.068 (3.861)***
ECM_{t-1}	-0.044 (-3.666)***	-0.299 (-4.054)***	-0.045 (-3.661)***	-0.299 (-4.054)***
R^2	0.494	0.517	0.494	0.517
F-stat.	9.381***	10.308***	9.370***	10.308***
DW	2.059	1.788	2.053	1.788
Hansen (1992b) stability tests for the ECMs				
Joint L_c	1.332 [0.290]	1.393 [0.240]	1.343 [0.280]	1.393 [0.240]
Error-variance	0.038 [0.940]	0.061 [0.800]	0.038 [0.940]	0.061 [0.800]

Note: see Table 1B.

Table 2A also reports parameter stability tests results. According to the Sup*F* and Mean*F* test results, the estimated models are not stable with regime shifts. However, the results for the L_C tests indicate that the long-run parameters are relatively constant through time, which also confirms that the $\ln(Y/L)$ series maintains a long-run level relationship with the $\ln(I/Y)$ series.

Table 2B presents the results for the short-run dynamic coefficients by estimating an ECM associated with the long-run estimates. The results indicate that the coefficients for the error correction terms (ECM_{t-1}) have the appropriate (negative) signs and are statistically significant at the 1% level. The estimated coefficient for ECM_{t-1} is −0.05 in the constant model and −0.30 in the trend model. To check the performance of the estimated error correction model, we also present some diagnostic tests associated with the model (R^2, *F*-statistic, Durbin-Watson (DW) test statistic for serial correlation).

Following Hansen (1992b), we report here the stability test results to check the stability of the error-variance and that of all parameters jointly. The joint stability test (L_C) results indicate that the null hypothesis of the parameter stability cannot be rejected at the 1% significance level in all specifications (Table 2B). Similar results are obtained for the error-variance at the 1% significance level. Coefficients of dummy variables are also significant when a linear deterministic term is included in the model.

In the short-run, the response of $\Delta\ln(Y/L)_t$ to $\Delta\ln(I/Y)_t$ does not depend on the selected ARDL model because its coefficient changes between 0.08 and 0.10. According to the results, a 1-point increase in the growth rate of the real investment share in real GDP leads to a 0.08- or 0.10-point increase in the growth rate of the real GDP per worker in the short-run.

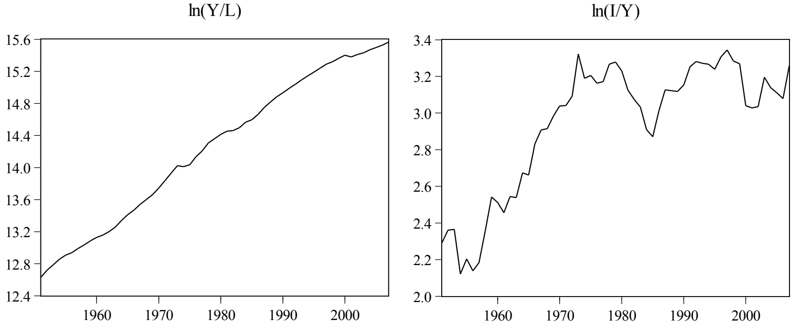
C. Empirical results for Taiwan

As an East Asian miracle economy (World Bank, 1993), the modern economic development of Taiwan can be divided into two distinct periods (Read, 2002). The first period corresponds to 1945-1960, and it refers to the inward-looking industrialisation period. During this inward-looking industrialisation period, the Taiwanese economy experienced strong GDP growth, 7.6% on average, between 1952 and 1960. The second period, beginning in 1960, marks the introduction of Taiwan's export-oriented growth strategy that has prevailed, with some modifications, to the present day.

As Wang (1997) notes, because Taiwan is a highly open economy, its development patterns and major fluctuations are closely related to international circumstances. Oil shocks in 1973 affected the Taiwanese economy severely. To reduce dependence on imported machinery equipment and intermediate inputs and to possibly increase the imports of raw materials, import-substitution with intermediate goods production was undertaken in the second half of the 1970s.

In the early 1980s, Taiwan began to run large and sustained trade surpluses, generating substantial foreign reserves. As a result of large trade surpluses with the United States, its key trading partner, Taiwan came under increasing external pressure to liberalise its economy, particularly its restrictions on imports, the inflows of foreign direct investment (FDI) and the under-valuation of the $NT. Domestic pressure for economic and political liberalisation was also increasing in the desire for both more market-oriented economic policies and institutional democratisation. The liberalisation process was initiated in the early 1980s with tariff reductions, eliminating some non-tariff barriers and removing restrictions on FDI in some previously reserved sectors. The economic and political reform processes then accelerated after the death of Chiang Ching-Kuo in January 1988. Despite a slow-down in the early 1980s, Taiwan continued to attain a high growth rate between 1980 and 1990, with the GDP growing by an average of 7.9% annually and with exports growing by an average of 9.7 percent annually (Read 2002).

Figure 3. Taiwan: real GDP per worker and investment share of real GDP



Note: the arguments of the log function are multiplied by 100.

Taiwan's growth performance between 1990 and 2000 remained buoyant despite representing a clear deceleration compared to the preceding three decades. Its annual GDP growth averaged 6.4%, and the $US value of exports averaged

9.9% during this period. This growth slow-down reflects the increasing maturity of Taiwan's economy and the sluggish regional growth in Asia since the 1997 crisis and the more recent global downturn. The Asian Crisis did not seriously affect Taiwan, although the $NT came under pressure from the ensuing financial downturn in the region. Taiwan's relatively robust economic structure and substantial foreign exchange reserves, in excess of $100 billion, provided a strong defence against the crisis, and growth in 1998 recovered relatively rapidly, though it was affected by the down-turn in regional economic activity (Read, 2002). When structural changes in the Taiwanese economy and the scatter diagram of $\ln(Y/L)$ and $\ln(I/Y)$ are combined, a dummy variable is defined covering the 1979-1997 period. $D7997$ takes a value of 1 between these years.

Table 3A gives the ARDL and FM-OLS estimates for the model. For Taiwan, AIC and SBC select different ARDL(p,q) models. The long-run parameters obtained from ARDL(5,0) and ARDL(2,0) are given in the first and fourth columns, and ARDL(1,0) and ARDL(1,0) are given in the second and fifth columns. Finally, the third and sixth columns of Table 3A give the FM-OLS estimates. The coefficient estimates indicate that the log of the investment share of the real GDP has a statistically significant positive effect on the log of real GDP per worker.

According to the results in Table 3A, when a linear deterministic trend term is added to the ARDL models, the long-run estimates of $\alpha/(1-\alpha)$ parameter are less than 1. The α parameter estimates thus change between 0.28 and 0.68 when a linear deterministic trend term is excluded from the model; only the L_C test shows that the estimated models are stable. The Mean$F$ and L_C tests cannot reject the null hypothesis that the long-run parameters are stable at the 1% level. The L_C statistics also confirm that the real GDP per worker equation level exists.

Table 3B reports the results of the short-run dynamic coefficients by estimating an ECM associated with the long-run estimates. The results indicate that the coefficients of the error correction terms (ECM_{t-1}) have negative signs and are statistically significant at the 1% level. The estimated ECM_{t-1} coefficient from different models changes between −0.035 and −0.139.

The joint stability test (L_C) results indicate that the null hypothesis of the parameter stability cannot be rejected at the 1% significance level in all specifications. Similar results are obtained for the error-variance at the 1% significance level. According to the estimation results, a 1-point increase in the growth rate of the investment share of real GDP leads to a 0.07- to 0.11-point increase in the growth rate of the real GDP per worker in the short-run. For Taiwan, there is evidence that the lagged changes of the $\Delta\ln(Y/L)$ series also affect the growth rate of the $\Delta\ln(Y/L)$ series. The D7997 coefficient is estimated to be statistically significant only when AIC chooses the appropriate lag and there is a trend term in the ARDL model.

Table 3. Taiwan

A. Long-run estimates of real GDP per worker equation with ARDL and FM-OLS

Dependent variable: ln(Y/L)	Without deterministic trend			With deterministic trend		
	AIC[a]	SBC[b]	FM-OLS	AIC[a]	SBC[b]	FM-OLS
Constant	10.033 (11.522)***	10.701 (9.115)***	10.166 (3.928)***	12.013 (46.169)***	11.965 (41.660)***	29.113 (5.864)***
Coefficient on ln(I/Y)	2.122 (7.537)***	1.858 (5.090)***	1.689 (1.938)*	0.391 (2.674)***	0.476 (2.678)***	0.930 (2.139)**
Estimate of α in eq. (6)	0.679	0.650	0.628	0.281	0.322	0.482
Stability tests of Hansen (1992a)						
Sup$F$		22.388			26.740	
Mean$F$		6.849			18.696♣	
L_c		0.556♣			0.687♣	

Note: see Table 1A. [a] The selected ARDL model is (5,0) and (2,0) by AIC. [b] The selected ARDL model is (1,0) and (1,0) by SBC.

B. Short-run estimates of real GDP per worker equation

Dependent variable: Δln(Y/L)	AIC: Without deterministic trend	AIC: With deterministic trend	SBC: Without deterministic trend	SBC: With deterministic trend
Intercept	0.002 (0.307)	0.006 (0.615)	0.002 (0.214)	0.006 (0.590)
$\Delta\ln(Y/L)_{t-1}$	0.137 (0.992)	0.252 (2.116)**	-	-
$\Delta\ln(Y/L)_{t-2}$	-0.271 (-1.992)*	-	-	-
$\Delta\ln(Y/L)_{t-3}$	-0.064 (-0.454)	-	-	-
$\Delta\ln(Y/L)_{t-4}$	-0.275 (-2.053)**	-	-	-
$\Delta\ln(I/Y)_t$	0.112 (3.415)***	0.074 (2.497)**	0.090 (2.890)***	0.083 (2.756)***
$D7997$	-0.006 (-0.409)	0.001 (0.109)*	-0.001 (-0.080)	0.005 (0.344)
ECM_{t-1}	-0.035 (-4.098)***	-0.139 (-3.391)***	-0.024 (-4.136)***	-0.130 (-4.445)***
R^2	0.363	0.372	0.315	0.340
F-stat.	2.354**	7.403***	7.974**	8.946***
DW	2.211	1.956	1.631	1.533
Hansen stability tests for the ECMs				
Joint L_c	1.331 [0.600]	0.721 [0.770]	0.589 [0.750]	0.682 [0.62]
Error-variance	0.114 [0.500]	0.169 [0.320]	0.198 [0.260]	0.255 [0.180]

Note: see Table 1B.

D. Growth accounting results

Table 4 reports growth accounting results. Our sources of growth calculations in Table 4 are based on short-term coefficients obtained from long-run information determined with ARDL and Harrod-neutral technological progress identification. Using this approach, we reconcile two propositions: growth accounting is a short-term issue and times series methods assume implicitly or explicitly that there are long-run equilibrium relationships and steady state conditions.

Our results justify the argument that the main source of economic growth is technological progress for the East Asian Tigers. Indeed, as it is shown in brackets in the last column of the Table 4, technological progress explains almost a hundred percent of the output per labour growth. The result derived here as the main source of economic growth is technological progress for the East Asian Tigers, is compatible to Klenow and Rodríguez-Clare (1997), who also assume the identifying assumption to be Harrod-neutral. Our results are compatible to Madsen (2010) which investigates the sources of growth for OECD countries, although Madsen (2010) assumes the identifying assumption to be Hicks-neutral. According to Madsen (2010: 756), the impact of total factor productivity (TFP) is magnified by a factor of $1/(1-\alpha)$ due to the endogeneity of capital deepening. For Madsen (2010), capital deepening has two effects. The first one is the direct effect that emphasises progress in production methods. The second effect is indirect and indicates a relationship between a higher TFP and an increase in expected earnings through a mechanism from the share market. Although Romer (1996: 26) indicates that short-run growth can result from either technological progress or capital accumulation, Madsen (2010) finds evidence that supports the following proposition: "In contrast to most growth accounting exercises, capital deepening was found to be an unimportant source of growth after taking into account that most capital deepening over the past 137 years has been TFP-induced" (Madsen 2010: 765).

Table 4. Growth accounting results based on the short-term coefficient from long-run information

Country	Average growth rate of output per labour (%) (1)	Contribution of capital stock per output (2)	Contribution of technological progress (3) =(1) −(2)
Hong-Kong	4.5199	-0.0012 [0.195] **	4.5211 (100.03)
		-0.0013 [0.209] ***	4.5212 (100.03)
Korea	3.8669	0.1643 [0.085] ***	3.7026 (95.75)
		0.1914 [0.099] ***	3.6755 (95.05)
		0.1952 [0.101] ***	3.6717 (94.95)
Taiwan	5.2389	0.1278 [0.074] **	5.1110 (97.56)
		0.1434 [0.083] ***	5.0955 (97.26)
		0.1554 [0.09] ***	5.0834 (97.03)
		0.1934 [0.112] ***	5.0454 (96.31)

Note: Numbers in parenthesis represent the percentage contributions. Numbers in brackets represent the short-term coefficient given in Tables 1B, 2B, and 3B. The symbols ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively. All of the results documented in Tables 2A and 3A are compatible with the assumption that the value of α is between zero and one. The results with deterministic trend documented in Table 1A are compatible to the assumption that the value of α is between zero and one. Source: authors' calculations.

VI. Conclusion

If the sources of economic growth are analysed based on time series econometrics, which generally examines long-run relationships among variables, one should assume that the nature of technological progress is Harrod-neutral. However, the sources of growth method assumes short-run relations. To solve this problem, the present study thus initially estimates the long-term coefficients based on a Harrod-neutral identifying assumption. It then uses the short-term coefficients yielded from the long-term relationship. The present study solves the contradiction between the econometric and economic theories for growth accounting.

The empirical findings indicate that the main source of economic growth is technological progress for Hong-Kong, the Republic of Korea, and Taiwan. We thus contribute to the sources of growth debate, which asks whether the sources of fast rates of growth in Hong Kong, the Republic of Korea, and Taiwan are capital accumulation or technological progress.[7]

Interestingly, approximately all per labour output growth stems from technological progress in the short-run. According to our identifying assumption, per labour output growth also equals the growth rate of technology in the long-run. The main source of economic growth is technological progress both in the short and long run.

We conclude by emphasising the harmony between theory and application. The theory proposes that the Solow model augmented with Harrod-neutral technological progress explains the long-run growth differences using the growth rate of technology. Our application results justify this theoretical point in the short-run. Productivity differences are the major reason explaining the output per labour growth differences among countries. However, one should recognize that it is assumed that human capital stock and all of the other components which have an impact on the level of technology or productivity are embodied in the labour force. According to our results, the main source of economic growth is not physical capital accumulation but human capital accumulation and all of the other variables which have an impact on the level of technology.

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[7] As noted in footnote 5, since all of the estimated values of α in equation (6) for Singapore are not compatible with the assumption that the value of α is between zero and one, the growth accounting results for Singapore are not discussed in this section.

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