

## **Does Financial Structure Matter?**

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### **Abstract**

We address the issue of whether financial structure influences economic growth. Three competing views of financial structure exist in the literature: the bank-based, the market-based and the financial services view. Recent empirical studies examine their relevance by utilising panel and cross-section approaches. This paper for the first time ever utilises time series data and methods, along with the Dynamic Heterogeneous Panel approach, essentially on developing countries. We find significant cross-country heterogeneity in the dynamics of financial structure and economic growth, and conclude that it is invalid to pool data across our sample countries. We find significant effects of financial structure on real per capita output, which is in sharp contrast to some of the recent findings. Panel estimates, in most cases, do not correspond to country specific estimates, and hence may proffer incorrect inferences for several countries of the panel.

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# Does Financial Structure Matter?

## 1. Introduction

Whether financial structure influences economic growth is a crucial policy issue. If one form of financial structure is more conducive to economic growth than another, then economic policy must take this into account. It is, therefore, hardly surprising that the distinction between bank-based and market-based financial systems, and their relative importance to economic growth, has been the focus of the theoretical debate for over a century (Allen and Gale, 1999; Gerschenkron, 1962; Stiglitz, 1985). The debate is still very alive (see, for example, Levine, 2002), not least because resolving this issue undoubtedly improves the quality of economic policies.

The empirical literature on this issue attempts to examine whether one type of financial system better explains economic growth than another. However, these studies are not without their own problems. The studies that analyse the UK and the US as market-based systems versus Japan and Germany as bank-based systems (e.g. Hoshi et al, 1991; Mork and Nakkamura, 1999; Weinstein and Yafeh, 1998; Arestis et al., 2001) tend to show that financial structure matters. However, they are susceptible to the criticism that these countries historically share similar growth rates. Therefore, they may not form a suitable sample to investigate the relative contribution of one financial system over another in the growth process (Goldsmith, 1969). Moreover, the results based on Japan, Germany, the UK and the US can only be used as a conjecture when it comes to economic policy for developing countries. Put it simply, the relationship between financial structure and economic growth remains unaddressed in the case of developing countries. We aim to fill this gap in this paper.

Panel and cross-section studies (Demirguc-Kunt and Levine, 1996; Levine, 2002 and 2003; Beck and Levine, 2002), find that financial structure is irrelevant to economic growth: neither the bank-based nor the market-based financial system can explain economic growth. Instead, it is the overall provision of financial services (banks and financial markets taken together) that are important. These multi-country studies are also subject to a number of concerns. Levine and Zervos (1996, p. 325) state that panel regressions mask important cross-country differences and suffer from 'measurement,

statistical, and conceptual' problems. Quah (1993; see, also, Caseli et al., 1996) shows the difficulties associated with the lack of balanced growth paths across countries when pooling data. Pesaran and Smith (1995) point out the heterogeneity of coefficients across countries. Luintel and Khan (2002) show that panel estimates often do not correspond to country-specific estimates. Consequently, generalisations based on panel results may proffer incorrect inferences for several countries of the panel. In short, panel estimates may be misleading at country level; consequently their policy relevance may be seriously impaired.

This paper contributes to the empirical literature surrounding financial structure and economic growth in a number of ways. First, we collect time series data on financial structure for six countries, most of which are developing economies.<sup>1</sup> Data span for a minimum of 30 (South Korea) to a maximum of 39 years (Greece). Our sample of low- and middle-income countries with varied growth experiences (see Table 1), also addresses the concern raised by Goldsmith (1969). Second, we formally test whether the data of our sample countries can be pooled. Both time series and dynamic heterogeneous panel methods, which do not impose any cross-country restrictions on parameters and adjustment dynamics, are applied. To our knowledge, this is the first ever study of this kind which estimates the long-run relationship between financial structure and economic growth using time series and dynamic heterogeneous panel methods, and utilising data from developing countries. Johansen's (1988, 1991) multivariate vector auto-regression (VAR), a well-established method in time series econometrics, is used. The robustness of our results is further checked through the Fully Modified OLS (FMOLS) estimators of Phillips and Hansen (1990). Panel estimates are obtained following the heterogeneous panel estimators due to Larson et al., (2001) and Pedroni (2001); so that (i) our results could be compared with those in the relevant literature (e.g. Levine, 2002; Beck and Levine, 2002); and (ii) we could formally test whether the panel estimates (parameters) correspond to the country specific estimates.

Our results are quite revealing. First, time series results show that for the majority of sample countries financial structure significantly explains economic growth. The

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<sup>1</sup> The six countries are: Greece, India, South Korea, the Philippines, South Africa and Taiwan.

results from the dynamic heterogeneous panels also confirm the significance of financial structure. These findings are in sharp contrast to existing results, which either depict financial structure as irrelevant (Levin, 2002 and Beck and Levine, 2002), or else that only bank-based financial systems are conducive to growth (Arestis, et al., 2001). Second, we find significant heterogeneity in cross-country parameters and adjustment dynamics; tests show that data cannot be pooled for these six countries. Tests also show that the panel parameters do not correspond to country specific estimates. Thus, our results uphold the assertion of Levine and Zervos (1996) that ‘panel regressions mask important cross-country differences’. Third, our results are robust to estimation methods and stability tests. Overall, our findings indicate that the apparent failure of large cross-country studies to identify a significant effect of financial structure on economic growth may be due to their failure to account sufficiently for cross-country heterogeneity.

The rest of the paper is organised as follows. In the section that follows we briefly discuss the theoretical arguments and the empirical evidence surrounding financial structure and economic growth. Section 3 outlines model specification and the econometric methods; section 4 discusses the dataset; section 5 covers tests of poolability; section 6 presents and discusses the main empirical results; and section 7 summarises and concludes.

## **2. Financial Structure and Development**

### *Theoretical Considerations*

The relationship between financial structure and economic development can be examined on the basis of competing theories of financial structure. These are: the bank-based, the market-based and the financial services. We discuss them briefly in what follows. The bank-based theory emphasises the positive role of banks in development and growth, and, also, stresses the shortcomings of market-based financial systems. It argues that banks can finance development more effectively than markets in developing economies, and, in the case of state-owned banks, market failures can be overcome and allocation of savings can be undertaken strategically (Gerschenkron, 1962). Those banks that are unhampered by regulatory restrictions, can exploit economies of scale and scope in information gathering and processing (for more details on these aspects of bank-based systems, see Levine, 2002, and Beck and

Levine, 2002). Indeed, bank-based financial systems are in a much better position than market-based systems to address agency problems and short-termism (Stiglitz, 1985; Singh, 1997). The bank-based view also stresses the shortcomings of market-based systems. The latter reveal information publicly, thereby reducing incentives for investors to seek and acquire information. Information asymmetries are thus accentuated, more so in market-based rather than in bank-based financial systems (Boyd and Prescott, 1986). Banks can ease distortions emanating from asymmetric information through forming long-run relationships with firms, and, through monitoring, contain moral hazard. As a result, bank-based arrangements can produce better improvement in resource allocation and corporate governance than market-based institutions (Stiglitz, 1985; Bhidé, 1993).

By contrast, the market-based theory highlights the advantages of well-functioning markets, and stresses the problems of bank-based financial systems. Big, liquid and well-functioning markets foster growth and profit incentives, enhance corporate governance and facilitate risk management (Levine, 2002, and Beck and Levine, 2002). The inherent inefficiencies of powerful banks are also stressed, for they “can stymie innovation by extracting informational rents and protecting firms with close bank-firm ties from competition ..... may collude with firm managers against other creditors and impede efficient corporate governance” (Levine, 2002, p. 3). Market-based financial systems reduce the inherent inefficiencies associated with banks and are, thus, better in enhancing economic development and growth. A related argument is that developed by Boyd and Smith (1998), who demonstrate through a model that allows for financial structure changes as countries go through different stages of development, that countries become more market-based as development proceeds. An issue of concern, identified by a recent World Bank (2001) study in the case of market-based financial systems in developing countries, is that of asymmetric information. It is argued that “the complexity of much of modern economic and business activity has greatly increased the variety of ways in which insiders can try to conceal firm performance. Although progress in technology, accounting, and legal practice has also improved the tools of detection, on balance the asymmetry of information between users and providers of funds has not been reduced as much in developing countries as it has in advanced economies – and indeed may have deteriorated” (p. 7).

The third theory, the financial services view (Merton and Bodie, 1995; Levine, 1997), is actually consistent with both the bank-based and the market-based views. Although it embraces both, it minimises their importance in the sense that the distinction between bank-based and market-based financial systems matters less than was previously thought; it is financial services themselves that are by far more important, than the form of their delivery (World Bank, 2001). In the financial services view, the issue is not the source of finance. It is rather the creation of an environment where financial services are soundly and efficiently provided. The emphasis is on the creation of better functioning banks and markets rather than on the type of financial structure. Quite simply, this theory suggests that it is neither banks nor markets that matter; it is both banks and markets. They are different components of the financial system; they do not compete, and as such ameliorate different costs, transaction and information, in the system (Boyd and Smith, 1998; Levine, 1997; Demirguc-Kunt and Levine, 2001). Under these circumstances, financial arrangements emerge to ameliorate market imperfections and provide financial services that are well placed to facilitate savings mobilisation and risk management, assess potential investment opportunities, exert corporate control, and enhance liquidity. Consequently, as Levine (2002) argues, “the financial services view places the analytical spotlight on how to create better functioning banks and markets, and relegates the bank-based versus market-based debate to the shadows” (p. 3).<sup>2</sup>

### *Empirical Evidence*

A number of studies have concentrated on comparisons that view Germany and Japan as bank-based systems, while the US and UK as market-based systems. These studies have employed rigorous country-specific measures of financial structure. Studies of

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<sup>2</sup> A special case of the financial services view is the law and finance view (La Porta et al, 1998; see, also, Levine, 1999). It maintains that the role of the legal system in creating a growth-promoting financial sector, with legal rights and enforcement mechanisms, facilitates both markets and intermediaries. It is, thereby, argued that this is by far a better way of studying financial systems rather than concentrating on bank-based or market-based systems. The World Bank (2001) view on the matter, based on “econometric results systematically points in one direction: far from impeding growth, better protection of the property rights of outside financiers favors financial market development and investment” (p. 8). Indeed, Rajan and Zingales (1998) argue that although countries with poor legal systems benefit from a bank-based system, better legal systems improve market-based systems, and as such the latter are preferable. While we recognise the importance of legal systems in a growth-promoting finance sector, we do not attempt to deal with this issue in this paper. It requires a study by itself and as such it is left for another occasion.

Germany and Japan use measures of whether banks own shares or whether a company has a 'main bank' respectively (Hoshi et al., 1991; Mork and Nakkamura, 1999; Weinstein and Yafeh, 1998). These studies provide evidence that confirms the distinction between bank-based and market-based financial systems in the case of the countries considered. However, reassessment of the evidence on the benefits of the Japanese financial system in view of the economy's poor performance in the 1990s has concluded against the beneficial effects of the bank-based nature of this system. Bank dependence can lead to a higher cost of funds for firms, since banks extract rent from their corporate customers (Weinstein and Yafeh, 1998). Studies of the US and the UK concentrate on the role of market takeovers as corporate control devices (Wenger and Kaserer, 1998; Levine, 1997), and conclude in favour of market-based financial systems. Goldsmith (1969), however, argues that such comparisons in the case of Germany and the UK for the period 1864-1914 does not contribute to the debate since "One cannot well claim that a superiority in the German financial structure was responsible for, or even contributed to, a more rapid growth of the German economy as a whole compared to the British economy in the half-century before World War I, since there was not significant difference in the rate of growth of the two economies" (p. 407). Our own study (Arestis et al., 2001) that provides evidence for the superiority of bank-based systems may be subjected to the same criticism. We may note in passing, though, that the implications for developing economies are evident, as argued in that paper.

Levine (2002) reinforces Goldsmith's (1969) argument when concluding that "financial structure did not matter much since the four countries have very similar long-run growth rates" (p. 4). Levine (op. cit.) addresses this problem by using a broad cross-country approach that allows treatment of financial system structure across many countries with different growth rates. The findings of this study support neither the bank-based nor the market-based views; they are, instead, supportive of the financial services view, that better-developed financial systems is what matters for economic growth. An earlier study by Demirguc-Kunt and Levine (1996), using data for forty-four industrial and developing countries for the period 1986 to 1993, had concluded that countries with well-developed market-based institutions also have well-developed bank-based institutions; and countries with weak market-based institutions also have weak bank-based institutions. Thereby supporting the view that

the distinction between bank-based and market-based financial systems is of no consequence. However, Levine and Zevros (1998), employing cross-country regressions for a number of countries covering the period 1976 to 1993, conclude that market-based systems provide different services from bank-based systems. In particular, market-based systems enhance growth through the provision of liquidity, which enables investment to be less risky, so that companies can have access to capital through liquid equity issues (see, also, Atje and Jovanovic, 1993, and Harris, 1997). The World Bank (2001) provides a comprehensive summary of the available evidence, which also reaches similar conclusions. It argues strongly that the evidence should be interpreted as clearly suggesting that “both development of banking and of market finance help economic growth: each can complement the other” (p. 48). In what follows we attempt to tackle the problem alluded to by Goldsmith (1969) and others. We also deal with the concerns surrounding the panel and cross-country regressions referred to by, among others, Levine and Zervos (1996). Our usage of time-series and heterogeneous panel estimators to analyse a number of diverse countries, should go some way in tackling these concerns.

### **3. Specification and Econometric Methods**

#### *Specification*

We specify a generalised Cobb-Douglas production function of the following form:

$$\log(Q/L)_t = a_0 + a_1 \log(K/L)_t + a_2 \log(\text{STR})_t \quad (1)$$

where Q is output, L is labour, K is capital and STR is financial structure (defined as the capitalisation ratio over bank lending ratio; see, also, below). Higher STR means a system that is more of the market-based variety; while a lower STR means more of a bank-based system. In specification (1), financial structure directly accounts for Total Factor Productivity (TFP). In actual estimations we use per capita output (LYP) and per capita capital stock (LKP), since consistent time series on labour force do not exist for most of the sample countries. It is important to note that for the purposes of this study, we are interested in the significance or otherwise of the coefficient  $a_2$ , rather than its sign. In either case a significant  $a_2$  coefficient implies that financial structure matters; an insignificant  $a_2$  coefficient implies that financial structure is of no consequence whatsoever.



It is common that cross-section studies use several other determinants of economic growth - the years of schooling (human capital), black market premiums, indicators of civil liberty, revolutions and coups, assassinations, bureaucratic efficiency, corruptions etc. However, data on these variables are usually obtained from periodic surveys and hence consistent time series are unavailable. Nevertheless, our specification (1) compares quite favourably with the ‘simple conditioning set’ specified by Levine (2002) and Beck and Levine (2002). They use initial levels of income and schooling as ‘simple conditioning set’ and examine the effect of financial structure on economic growth in panel/cross sectional framework, whereas we specify a generalised Cobb-Douglas production function.<sup>3</sup>

### *Econometric Methodology*

Under the Johansen (1988) maximum likelihood (ML) approach a k-dimensional and p<sup>th</sup> order vector (X) can be re-parameterised to a vector error-correction model (VECM):

$$\Delta X_t = \mu + \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + \dots + \Gamma_{p-1} \Delta X_{t-p+1} + \Pi X_{t-p} + \varphi D_t + \varepsilon_t \quad (2)$$

In our analysis  $X_t$  [LYP, LKP, STR] is a 3x1 vector of first-order integrated [I(1)] variables;  $\Gamma_i$  are (3x3) short-run coefficient matrices;  $\Pi_{(3 \times 3)}$  is a matrix of long-run (level) parameters;  $D_t$  captures the usual deterministic components;  $\mu$  is a constant term and  $\varepsilon_t$  is a vector of Gaussian error. A co-integrated system,  $X_t$ , implies that: (i)  $\Pi = \alpha_{(3 \times r)} \beta'_{(r \times 3)}$  is rank deficient, i.e.  $r < k$  ( $r$  = number of distinct cointegrating vectors;  $k = 3$ ); and (ii)  $\{\alpha_{\perp} \Gamma \beta_{\perp}\}$  has full rank,  $(k-r)$ , where  $\alpha_{\perp}$  and  $\beta_{\perp}$  are  $(3 \times (3-r))$  orthogonal matrices to  $\alpha$  and  $\beta$ . The rank of  $\Pi$  is tested by the well known Maximal Eigenvalue ( $\lambda$ -max) and Trace statistics (Johansen, 1988).

A number of issues are important for the estimation of VAR models. It is the time span of the data rather than the number of observations, which determines the power of cointegration tests (Campbell and Perron, 1991). Our data extend from a minimum of 30 (South Korea) to a maximum of 39 (Greece) years, which in our view, provides

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<sup>3</sup> Barro and Lee (2000) and Cohen and Soto (2001) provide periodic data, five yearly and 10 yearly respectively, on educational attainment for several countries of the world. We thought of interpolating annual series on educational attainments for our sample countries from these periodic observations. However, the owners of the respective data sets advised us strongly against interpolation, on the grounds of unreliability. We, therefore, decided not to pursue this matter any further.

sufficient time length to capture the long-run relationship between LYP, LKP and STR. We specify VAR lengths ( $p$ ) such that the VAR residuals are rendered non-autocorrelated.<sup>4</sup> A constant term is entered in the co-integrating space to allow for non-zero mean of the system variables. A trivariate VAR can exhibit two cointegrating vectors at the most. Pesaran and Shin (2002) suggest identification of multi-cointegration through the tests of over-identifying restrictions. We follow their approach of identification if multiple cointegrating vectors are found.

We examine the robustness of our VAR results by employing the FMOLS estimators. The Johansen method is a reduced-form dynamic system estimator and addresses the issues of multi-cointegration and normalisation. The FMOLS estimator of Phillips and Hansen (1990), on the other hand, is a single equation estimator of long-run parameters when variables are  $I(1)$ . FMOLS corrects for both short- and long-run dependence across equation errors, and it is shown to be super-consistent, asymptotically unbiased and normally distributed. The associated t-ratios allow inference using standard tables. In short, FMOLS corrects the bias and non-normality inherent in the Engle-Granger (1987) OLS estimator of cointegrating parameters. This robustness check is important particularly in view of the different formulations of co-integration tests between these two estimators. It also indicates the sensitivity of our results to estimation methods. In the event of contradictory results, we attach more weight to the results based on the system estimator. A brief outline of the FMOLS estimator is provided in the Appendix.

#### **4. Description of Data**

Our sample consists of six countries, viz. Greece, India, South Korea, the Philippines, South Africa and Taiwan. Data on real Gross Domestic Product (GDP), real gross fixed Investment (I), Bank Lending Ratio (BLR, defined as total lending by deposit taking institutions/GDP) and total population are obtained from IMF CD-ROM (March 2002). Capitalization Ratios (CLR, defined as total value of domestic equities listed in domestic stock exchanges/GDP) are obtained from Global Financial Data,

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<sup>4</sup> Johansen (1992) suggests that the lag length in the VAR should be specified whereby the VAR residuals are rendered uncorrelated. Selection of lag length based on information criteria may not be adequate to render the VAR residual uncorrelated (Cheung and Lai, 1993). Hence, we specify lag-length based on the test of serial correlation in VAR residuals.

Inc. and Standard and Poor's Emerging Markets database (2002). Data frequency, determined simply by data availability, is annual and the sample period is 1962-2000 for Greece, 1970-1999 for South Korea, 1966-1999 for India, 1969-1999 for the Philippines, 1965-1999 for South Africa and 1965-2000 for Taiwan. The heterogeneity in the sample period across countries is due to unavailability of data. A consistent series of total physical capital stock for the whole sample period is, unfortunately, not available. Therefore, we constructed it for each country in the sample from the respective real gross fixed investment series using the perpetual inventory method. Following Luintel and Khan (1999), amongst others, a depreciation rate of eight percent and the sample-average growth rate of real investment, are used to compute the initial capital stock. Following Levin (2002) and Beck and Levine (2002), financial structure is defined as the log of the capitalisation ratio over the bank-lending ratio. Thus, our measure of financial structure is akin to their measure of 'structure-size'.

Table 1 near here

Table 1 reports some descriptive statistics of our data set. It is obvious that our sample consists of countries with differing income levels and varied growth experience. In our sample Korea was the fastest growing economy (7.6 % growth of real per capita income per annum) and the Philippines the slowest (0.6% per capita income growth per annum). A striking feature, however, is that all but one (Taiwan) sample countries have evolved towards a more market-based system over the last thirty to forty years. Although the bank lending ratios have gone up for all sample countries during this period, the rise in capitalisation ratio is by far greater. The bank-lending ratio, on average, is 2.73 folds higher in the last five years of the sample compared to its level of first five year, but the capitalisation ratio has shot up by 7.60 folds during the same period. The low base may partly explain this huge rise in the capitalisation ratio. Nevertheless, financial structure, on average, has gone up by almost three folds in the intervening period. The last column of Table 1 shows that the financial systems of sample countries grew towards a more market-oriented system by an average annual rate of 0.2 % for Taiwan (lowest) and 2.5% for Korea (highest). Figure 1 plots LYP for all sample countries.

Figure 1 near here

All plots are normalised at 1995=1 for the ease of comparison across countries. In econometric modelling we do not use normalised data. It is apparent that LYP shows positive trend for all countries but at varying rate. Taiwan and South Korea show quite steep rise in real per capita output whereas plots for India and the Philippines appear relatively flat. Plots for Greece and South Africa are in between.

**Figure 2 near here**

Figure 2 plots LKP. Again, the rate of capital accumulation appears quite high for Taiwan and South Korea, as their plots of LKP are pretty steep. Greece shows a rapid rate of capital accumulation prior to 1975, but it slows down thereafter. For the remaining countries the rate of accumulation process appears rather slow as depicted by the flatness of their plots.

**Figure 3 near here**

Figure 3 plots the financial structure variable. Plots appear more volatile than those of LYP and LKP; a positive trend is also not easily discernable in some cases.<sup>5</sup> Greek financial structure depicts big spikes during the 1970s and late 1990s. Taiwan and South Korea also show spikes during late 1980s but of a lesser magnitude. Overall, these plots show a gradual move towards a market oriented financial system.

## **5. Heterogeneity**

Our sample consists of low- and middle-income countries, which represent different stage of development and economic structure. They also share significantly different growth experiences (Table 1). It is, therefore, interesting to formally test if it is valid to pool the data set of these countries. This is important not least because there is a growing concern about the panel and cross-section tests, in that they neglect heterogeneity.

Formal tests of the dynamic heterogeneity of financial structure and economic growth are conducted as follows. First, we estimate a series of  $p^{\text{th}}$  ( $p=1,2,3$ ) order autoregressive and distributed lag models, ADL(P), conditioning LYP on LKP and STR, and test for the equality of parameters across sample countries. Second, we estimate ADL(P) on growth rates and perform tests of parameter equality.

**Table 2 near here**

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<sup>5</sup> This lack of apparent positive trend in some of these plots is mainly due to the big scaling on the vertical axis, which is required to accommodate the Geek series. Country-by-country plot shows a positive trend more clearly.

Chow-type F tests under the null of parameter equality across sample countries are reported in Table 2, where the tests reject the null under all specifications. Thus, the elasticity of LYP with respect to LKP and STR is heterogeneous across countries. Furthermore, as another measure of dynamic heterogeneity, we test for error variances homoskedasticity across groups. The LM-test of group-wise heteroskedasticity is reported in Table 2, which confirms that error variances across sample countries are significantly different and this also holds across all specifications. It follows that the elasticity of LYP with respect to LKP and STR, as well as the error dynamics across sample countries, are significantly heterogeneous. Consequently, the data set cannot be pooled. This raises concerns with respect to the validity of extant panel, and cross-sectional tests that do not allow for cross-country heterogeneity.

## 6. Empirical Results

In order to evaluate the time series properties of the data formally, we implement the univariate KPSS test (Kwiatkowski et al., 1992), which tests the null of stationarity.<sup>6</sup> The results are reported in Table 3.

Table 3 near here

LYP and LKP are non-stationary; tests reject the null of stationarity in all cases but one, this being the Philippines' LYP. The latter's trend stationarity is rejected but level stationarity is not. The financial structure variable also appears non-stationary in all but two cases. The exceptions are Greece and South Korea. The Greek financial structure appears level stationary but not trend stationary whereas the opposite holds for South Korea. We further examine the autocorrelation functions for these (three) suspects and found that they decay slowly which means they appear closer to I(1) series than to I(0). Hence, we treat them as I(1) in further modelling. All series appear unequivocally stationary in their first differences. Thus, the overall finding of the KPSS tests is that LYP, LKP and STR are I(1).

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<sup>6</sup> Kwiatkowski et al. (1992) show that these tests are more powerful than the usual DF/ADF tests. Recently, however, Caner and Kilian (2001) warn against these power gains, especially for high frequency data. Our data are low frequency.

Table 4 near here

Table 4 reports the Johansen rank tests and a range of VAR diagnostics obtained from the VECM. Trace tests show that LYP, LKP and STR are co-integrated and exhibit a single co-integrating rank (vector) for all sample countries. The  $\lambda$ -max statistic supports these findings except that the evidence of co-integration for India and the Philippines now appears marginal. Given the superiority of trace statistics over the maximal eigen-value statistics in testing the null of non-co-integration, we conclude that LYP, LKP and STR are co-integrated with single co-integrating vector for all of our sample countries.

For a valid normalisation and error-correction representation, the associated loading factors ( $\alpha_s$ ) must be negatively signed and significant. On this basis, we can normalise all countries on LYP; their associated loading factors are negatively signed and significant at 5% or better. LM tests show absence of serial correlation in VAR residuals for all cases. The VAR residuals pass normality tests except for Taiwan. Thus, utilizing the VECM, we are able to identify a long-run output relationship based on per capita capital stock and the financial structure variable, which confirms the error-correcting behaviour when displaced from the long-run equilibrium. Our empirical model also passes all the relevant diagnostics.

Table 5 near here

Table 5 (section A) reports the co-integrating vectors (long-run parameters) based on both the Johansen VECM and the FMOLS estimator. In the last column of the Table we report the tests of stationarity of the error-correction term derived from the FMOLS estimates, which is equivalent to the tests of co-integration. FMOLS corroborates our earlier findings of VECM that the series in question (LYP, LKP and STR) are co-integrated for all sample countries.

As expected the long-run elasticity of LYP with respect to LKP is positive and highly significant for all the countries. This holds true under both estimators. However, the elasticity of LYP with respect to STR is not as universal. Under the Johansen approach, financial structure significantly affects per capita GDP for all but one country (the Philippines). The sign and the significance of parameters show that the market-based financial system appears conducive to Greece, India, South Korea and

Taiwan whereas the bank-based system appears better for South Africa in explaining the long-run per capita output. In the case of the Philippines, financial structure appears insignificant in explaining per capita output. The parameters obtained from FMOLS are qualitatively similar to those obtained from VECM with only one exception: financial structure appears insignificant for South Africa. However, and as stated above, we attach more importance to the VECM results. Overall, we find significant effects of financial structure on long run per capita output in the majority of cases examined (five out of six countries). Indeed, our results are robust to the estimation methods.

In section B of Table 5 we report the panel, ‘between-dimension’, estimates of the elasticities of LYP with respect to LKP and STR. Larsson et al. (2001) discuss the computations of these panel estimates under the Johansen approach and Pedroni (2001) derives them for the FMOLS. Essentially, the ‘between-dimension’ panel parameters are the mean of the country-specific parameters. Our panel-based results corroborate our time series findings: financial structure appears significant for the panel of six countries, and this holds under both estimators (Johansen and FMOLS). This is in contrast to the findings of Levine (2002) and Beck and Levine (2002).<sup>7</sup> However, the signs of the coefficients are different across two estimators. VECM shows the bank-based system to be more conducive to the panel of six countries, whereas the FMOLS shows quite the opposite. It is important to note that under VECM, the negatively signed large coefficient (-0.519) of South Africa alone is sufficient to turn the overall coefficient for the panel into negative (-0.008). All the existing panel tests suffer from this typical caveat – results of one or few countries dominate the whole panel – and one of the contributions of our results is that they

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<sup>7</sup> It should be noted that the ‘between-dimension’ panel approach we use and those of Levin (2002) and Beck and Levine (2002) are not identical. These ‘between-dimension’ approaches allow for the cross-country heterogeneity whereas the approaches of Levin (2002) and Beck and Levine (2002) do not allow for cross-country heterogeneity. Nevertheless, comparison is ventured on the grounds that both are panel-based tests. As a matter of fact, the cointegration-based ‘between-dimension’ panel test (which we implement) is a statistically superior test than those implemented by Levine (op. cit.) and Beck and Levine (op. cit.).

bring this issue to focus.<sup>8</sup> This further lends support to our preference to country-by-country (time series-based) results.

Table 6 near here

The magnitude of the country-specific point estimates (elasticities) in Table 5 show a considerable degree of cross-country heterogeneity. From a policy perspective, it is extremely important to establish the degree of equivalence between the panel and the country specific estimates because national policies rely on such key parameters. We address this issue by formally testing if country-specific parameters are jointly equal to the corresponding panel estimates. This involves conducting a Wald or LR test for the restriction that each country-specific coefficient is equal to its panel counterpart and summing up the individual  $\chi^2$  statistics (Pesaran et al., 2000). Assuming that these tests are independent across countries, the sum of the individual  $\chi^2$  statistics tests the null that country-specific coefficients are jointly equal to their respective panel estimates. The test statistic is  $\chi^2(N)$  distributed, where N is the number of countries in the panel. Results of these joint tests are reported in Table 6 and they strongly reject the null of parameter equality except for the LKP under the Johansen approach.

Overall, we find that financial structure exhibits a significant effect on the level of output for most sample countries (five of the six) and results appear robust to the estimation methods utilised. The cross-country heterogeneity in estimated point elasticities are pervasive and statistically significant. Tests show that in most cases panel estimates do not correspond to country specific estimates and this heterogeneity appears more serious with respect to financial structure.<sup>9</sup>

### *Stability*

It is well known that structural shifts should be identified endogenously rather than exogenously (see, among others, Perron, 1997; Christiano, 1992; Quintos, 1995;

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<sup>8</sup> Panel unit root tests, panel co-integration tests (dynamic heterogeneous or otherwise) and traditional (OLS- and/or IV- based) panel tests all suffer from this problem.

<sup>9</sup> At country level, the null of equality of panel and country-specific elasticity of LYP with respect to STR is rejected in all but one case (the Philippines) under the Johansen estimator whereas FMOLS shows relatively few rejections. With respect to LKP, FMOLS shows more rejections (in four out of six sample countries) than the Johansen approach. Overall, country-by-country tests support the joint tests, that panel estimates do not correspond to the country specific parameters.



Luintel, 2000). In view of this proposition and our preference for VECM, we follow the recursive approach of Hansen and Johansen (1999). This approach essentially compares the recursively-computed ranks of the  $\Pi$  matrix with its full sample rank. A significant difference between them implies a structural shift in the cointegrating rank. Likewise, conditional on the identified ranks of  $\Pi$ , if sub-sample parameters significantly differ from those of the full sample, this signifies instability of the cointegrating parameters. The LR test for these hypotheses is asymptotically  $\chi^2$ , with  $kr-r^2$  degrees of freedom. Tests are carried out in two settings: (i) allowing both short-run and long-run parameters to vary (the Z-model); and (ii) short-run parameters are fixed and only long-run parameters are allowed to vary (the R-model).

We specify a base estimation window of the first 20 observations.<sup>10</sup> Since sample sizes differ across countries, the period of stability tests is not uniform – the stability tests range between a minimum of 13 years (South Korea) to a maximum of 17 years (Greece). Figure 4 plots the normalised LR statistics that test rank stability using the R-model.<sup>11</sup> All LR statistics are scaled by the 5% critical value; hence, values greater than unity imply rejection of the null of stability and vice versa. In these plots stability of rank,  $r$ , requires rejection of  $r-1$  ranks.

Figure 4 near here

Plots of the scaled LR statistics show that the null of non-cointegration ( $H_0: r=0$ ) is clearly rejected for all sample countries. All plots that test the null of  $r=0$  cross the critical threshold. The plots that test  $H_0: r \leq 1$  are all below unity (i.e. less than the 5% critical value) except for a marginal break shown by India during 1994-95. Figure 5 plots the normalised LR statistics, which test for the stability of cointegrating parameters.

Figure 5 near here

Plots pertaining to both Z- and R-models are reported. Co-integrating parameters are stable for all countries with only one exception. The Z- model shows parameter instability for South Africa prior to 1990; this is primarily due to the volatility of

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<sup>10</sup> Hansen and Johansen (1999) specify an initial estimation window of 16 (monthly) observations.

<sup>11</sup> The R-model is more suitable for testing the stability of cointegrating ranks and long-run parameters (Hansen and Johansen, 1999). Nonetheless, results from the Z-model appear broadly similar and hence are not reported (but are available from the authors upon request).

short-run parameters since the R-model shows parameter stability. Overall, our estimated co-integrating rank and parameters are remarkably stable.<sup>12</sup>

## **7. Summary and Conclusions**

We have focused in this paper on the important, but controversial, issue of whether financial structure matters in an economic system. We briefly reviewed the relevant theoretical and empirical literature before embarking on a time-series investigation of this issue, the first of this kind to be investigated. We further computed panel results based on dynamic heterogeneous panel estimators.

Our results clearly show that significant cross-country heterogeneity exists in financial structure and growth dynamics and it is invalid to pool data even for these six countries. This indicates that extant panel and/or cross-section studies of financial structure and economic growth, which pool several countries, may well have concealed important cross-country differences. We find a robust co-integrating relationship between output per capita, capital stock per capita and the financial structure. Financial structure exerts significant effects on the level of output per capita in all but one country (the Philippines). Furthermore, the magnitude of the long-run effects (cointegrating parameter) of financial structure on per capita output is extremely heterogeneous across countries. Tests reject the null of equality between the ‘between-dimension’ panel and country specific parameters in all cases but one (the elasticities of LKP under the Johansen method). Thus, panel estimates do not appear to correspond to country specific estimates (parameters). The speed of adjustment to long-run disequilibria also differs significantly across countries. A comparison of our time series and panel results also reveals that a single country may sufficiently dominate the result for the whole panel.

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<sup>12</sup> Tests of the stability of cointegrating parameters are also conducted under FMOLS by computing recursive Wald tests for the coefficients of STR and LKP. The null hypotheses are that the recursively computed sub-sample and full-sample parameters are equal. In the cases of South Korea and the Philippines they do not reject the null of stability. The other four countries show instances of instability and the coefficient of STR is found to be relatively more stable than that of the LKP. Overall, the results appear to corroborate the findings of system estimator that the elasticity of LYP with respect to STR is stable.

Our findings of a significant effect of financial structure on output levels are in sharp contrast to those of Levine (2002) and Beck and Levine (2002), amongst others. This contrast is maintained by both the empirical approaches (time series and the dynamic heterogeneous panel estimators) we have pursued in this study. It is, thus, possible that the apparent insignificant effect of financial structure on growth shown by extant panel tests may be due to their failure to address cross-country heterogeneity.

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## **APPENDIX**

### **The FMOLS Estimator**

Consider the following linear static regression:

$$y_t = \beta_0 + \beta_1' x_t + u_t \quad (\text{A1})$$

where  $y_t$  is a vector of I(1) dependent variable and  $x_t$  is (kx1) vector of I(1) regressors. Define  $\Delta x_t = \mu + w_t$ , where  $\mu$  is a (kx1) vector of drift parameters and  $w_t$  is

a  $(k \times 1)$  vector of stationary variables. Let  $\widehat{\xi} = (\widehat{u}_t, \widehat{w}_t)'$ , where a hat indicates a consistent estimator. Define the long-run variance-covariance matrix of  $\widehat{\xi}$  ( $\widehat{V}$ ):

$$\widehat{V} = \widehat{\Gamma} + \widehat{\Phi} + \widehat{\Phi}' = \begin{bmatrix} \widehat{v}_{11} & \widehat{v}_{12} \\ \widehat{v}_{21} & \widehat{v}_{22} \end{bmatrix} \quad (\text{A2})$$

and, further define,

$$\widehat{\Delta} = \widehat{\Gamma} + \widehat{\Phi} = \begin{bmatrix} \widehat{\Delta}_{11} & \widehat{\Delta}_{12} \\ \widehat{\Delta}_{21} & \widehat{\Delta}_{22} \end{bmatrix} \quad (\text{A3})$$

$$\widehat{Z} = \widehat{\Delta}_{21} - \widehat{\Delta}_{22} \widehat{v}_{22}^{-1} \widehat{v}_{21} \quad (\text{A4})$$

where  $\widehat{\Gamma} = \frac{1}{T-1} \sum_{t=2}^T \widehat{\xi}_t \widehat{\xi}_t'$ ;  $\widehat{\Phi} = \sum_{s=1}^m w(s, m) \widehat{\Gamma}_s$ ;  $\widehat{\Gamma}_s = T^{-1} \sum_{t=1}^{t-s} \widehat{\xi}_t \widehat{\xi}_{t+s}'$ ; and  $w(s, m)$  is the lag

truncation window. The adjustment in  $y_t$  is achieved by:  $\widehat{y}_t^* = y_t - \widehat{v}_{12} \widehat{v}_{22}^{-1} \widehat{w}_t$ . The

FMOLS estimator is:

$$\widehat{\beta}_{fmols} = (W'W)^{-1} (W' \widehat{y}^*) - TD\widehat{Z} \quad (\text{A5})$$

where  $\widehat{y}^* = (\widehat{y}_1^*, \widehat{y}_2^*, \dots, \widehat{y}_t^*)'$ ;  $D = [0_{1 \times k} \quad I_k]'$ ; and  $W_{(txk)}$  is matrix of regressors including a constant term. A consistent estimator of the variance-covariance matrix ( $\Psi$ ) is:  $\Psi(\widehat{\beta}_{fmols}) = \kappa_{11.2} (W'W)^{-1}$ , where  $\kappa_{11.2} = \widehat{v}_{11} - \widehat{v}_{12} \widehat{v}_{22}^{-1} \widehat{v}_{21}$ . The stationarity of the error correction term generated through  $\widehat{\beta}_{fmols}$  implies that (A1) is a co-integrated relationship.

**Table 1: Some Summary Statistics of Data**

	Growth of Real Per Capita Income	Bank Lending Ratio (BLR)				Capitalisation Ratio (CLR)				Financial Structure (STR)			
		BLR <sub>0</sub>	BLR <sub>T</sub>	BLR <sub>μ</sub>	ΔBLR	CLR <sub>0</sub>	CLR <sub>T</sub>	CLR <sub>μ</sub>	ΔCLR	STR <sub>0</sub>	STR <sub>T</sub>	STR <sub>μ</sub>	ΔSTR
Greece	0.037	0.273	0.921	0.720	0.020	0.043	0.757	0.198	0.024	0.156	0.807	0.302	0.020
India	0.026	0.248	0.462	0.409	0.007	0.036	0.337	0.113	0.007	0.146	0.729	0.254	0.013
South Korea	0.076	0.393	0.703	0.520	0.017	0.066	0.368	0.224	0.024	0.167	0.516	0.401	0.025
Philippines	0.006	0.281	0.671	0.384	0.011	0.119	0.674	0.254	0.011	0.423	1.030	0.632	0.005
South Africa	0.015	0.900	1.431	1.026	0.019	0.876	1.549	1.043	0.022	0.976	1.086	1.010	0.004
Taiwan	0.064	0.346	1.867	0.978	0.045	0.183	1.033	0.417	0.020	0.550	0.554	0.373	0.002
Average Change	0.037	2.733				7.603				2.968			

BLR = total lending by deposit taking institutions/GDP; CLR = total value of domestic equities listed in domestic stock exchange/GDP; STR = log(CLR/BLR). Growth of Real Per Capita Income indicates the average annual growth rate (expressed as ratio) over the sample period. Subscripts 0 and T denote mean values of the first five years and the last five years of the sample for each country; subscript μ indicates the sample mean value. The average change (in the last row) is calculated as:  $\sum_1^6 (X_t - X_0) / 6$ , where X denotes BLR, CLR and STR.



**Table 2: Tests of heterogeneity of financial structure and growth dynamics across sample countries**

	Specification: A			Specification: B		
	P=1	P=2	P=3	P=1	P=2	P=3
Parameter equality	17.352 <sup>a</sup> (5, 171)	19.299 <sup>a</sup> (7, 153)	12.083 <sup>a</sup> (10, 129)	9.389 <sup>a</sup> (5, 165)	8.577 <sup>a</sup> (7, 147)	5.833 <sup>a</sup> (10, 123)
LM Test	14.508 <sup>b</sup>	15.450 <sup>a</sup>	12.492 <sup>b</sup>	11.169 <sup>b</sup>	11.115 <sup>b</sup>	12.551 <sup>b</sup>

The specification A:  $LYP = \lambda_0 + \sum_{i=1}^p \lambda_{1i} LYP_{t-i} + \sum_{i=1}^p \lambda_{2i} LKP_{t-i} + \sum_{i=1}^p \lambda_{3i} STR_{t-i} + \varepsilon_t$ .

The specification B:  $\Delta LYP = \theta_0 + \sum_{i=1}^p \theta_{1i} \Delta LYP_{t-i} + \sum_{i=1}^p \theta_{2i} \Delta LKP_{t-i} + \sum_{i=1}^p \theta_{3i} \Delta STR_{t-i} + \varepsilon_t$ .

Equality of  $\theta$  and  $\lambda$  are standard (Chow type) F-tests of parameter equality across the sample (six) countries. Numbers within parentheses, (.), are the degrees of freedom of F distribution. The 1% critical value for F(5, 125) is 3.17; the 5% critical value for F(5, 125) is 2.29. Lagrange Multiplier (LM) tests (see text) reject the null of homoskedastic error variances across the sample countries; they are  $\chi^2(5)$  distributed. Superscripts 'a' and 'b' indicate rejection of the null at 1% and 5%. Variable definitions are: LYP = log of per capita real GDP; LKP = log of per capita real physical capital stock; and STR = log (CLR/BLR).

**Table 3: KPSS unit root tests**

Countries	LYP		LKP		STR	
	$\eta_{\mu}$	$\iota_{\mu}$	$\eta_{\mu}$	$\iota_{\mu}$	$\eta_{\mu}$	$\iota_{\mu}$
Greece	1.256 <sup>a</sup>	0.2338 <sup>a</sup>	1.254 <sup>a</sup>	0.326 <sup>a</sup>	0.207	0.198 <sup>b</sup>
India	1.151 <sup>a</sup>	0.248 <sup>a</sup>	1.182 <sup>a</sup>	0.292 <sup>a</sup>	0.704 <sup>b</sup>	0.278 <sup>a</sup>
South Korea	1.079 <sup>a</sup>	0.132 <sup>c</sup>	1.086 <sup>a</sup>	0.134 <sup>c</sup>	0.642 <sup>b</sup>	0.078
Philippines	0.312	0.136 <sup>c</sup>	0.893 <sup>a</sup>	0.259 <sup>a</sup>	0.452 <sup>c</sup>	0.186 <sup>b</sup>
South Africa	0.994 <sup>a</sup>	0.266 <sup>a</sup>	0.866 <sup>a</sup>	0.319 <sup>b</sup>	0.505 <sup>b</sup>	0.116 <sup>c</sup>
Taiwan	1.393 <sup>a</sup>	0.204 <sup>b</sup>	1.379 <sup>a</sup>	0.291 <sup>a</sup>	0.368 <sup>c</sup>	0.194 <sup>b</sup>

The critical values for  $\eta_{\mu}$  are 0.739, 0.463 and 0.347 at 1%, 5% and 10%; the respective critical values for  $\iota_{\mu}$  are 0.216, 0.146 and 0.119.  $\eta_{\mu}$  and  $\iota_{\mu}$  respectively test the nulls of level and trend stationarity. In their first differences all series are stationary. The latter set of results is not reported to conserve space; they are available from the authors upon request. Superscripts a, b, and c indicate rejection of the null of stationarity at 1%, 5% and 10%, respectively. Variable definitions are: LYP = log of per capita real GDP; LKP = log of per capita real physical capital stock; and STR = log(CLR/BLR).

**Table 4: Co-integration tests and VAR diagnostics between LYP, LKP and STR (Johansen Method)**

	Trace Statistics, $H_0$ :			Maximal Eigenvalue, $H_0$ :			Loading Factor ( $\alpha$ )	LM{3}	NOR	LAG
	$r=0$	$r \leq 1$	$r \leq 2$	$r=0$	$r \leq 1$	$r \leq 2$				
GRE	39.53 <sup>b</sup> (0.015)	15.63 (0.196)	4.19 (0.396)	23.89 <sup>b</sup> (0.027)	11.44 (0.228)	4.19 (0.396)	-0.401 <sup>a</sup> [0.078]	0.748	0.192	2
IND	34.53 <sup>b</sup> (0.057)	14.85 (0.241)	3.00 (0.589)	19.68 (0.113)	11.85 (0.201)	3.00 (0.588)	-0.919 <sup>a</sup> [0.248]	0.188	0.334	2
KOR	32.60 <sup>c</sup> (0.092)	10.03 (0.643)	2.98 (0.593)	22.58 <sup>b</sup> (0.043)	7.05 (0.669)	2.98 (0.592)	-0.404 [0.123]	0.453	0.128	3
PHL	36.24 <sup>b</sup> (0.037)	17.46 (0.117)	6.40 (0.168)	18.78 (0.148)	11.07 (0.256)	6.40 (0.167)	-0.652 <sup>b</sup> [0.296]	0.795	0.115	2
SAF	34.92 <sup>b</sup> (0.052)	11.79 (0.477)	2.36 (0.707)	23.13 <sup>b</sup> (0.036)	9.43 (0.402)	2.36 (0.706)	-0.185 <sup>a</sup> [0.046]	0.114	0.215	2
TWN	44.40 <sup>a</sup> (0.003)	17.86 (0.110)	7.22 (0.120)	26.53 <sup>a</sup> (0.010)	10.64 (0.290)	7.22 (0.120)	-0.683 <sup>a</sup> [0.138]	0.146	0.000	3

The country mnemonics are: GRE=Greece; IND=India; KOR= South Korea; PHL= the Philippines; SAF= South Africa; TWN=Taiwan. Figures within parenthesis (.) are p-values under the  $H_0$ :  $r=0$ ;  $r \leq 1$  and  $r \leq 2$ . For the loading factors ( $\alpha$ ) figures within the brackets [.] are standard errors. LM{3} reports p-values of the third order LM test of the null of no serial correlation in VAR residuals. The column NOR reports p-values of Bera-Jarque normality tests of VAR residuals,  $\chi^2(2)$  distributed. The column LAG reports the VAR lag lengths used. Superscript a, b, and c indicates significance at 1%, 5% and 10% respectively. GRE, IND and PHL do not require any dummy. KOR required impulse dummies for 1978 and 1998; Taiwan required an impulse dummy around first oil price shock (1970 and 1971), although non-normality is still evident. SAF required impulse dummies for 1981 and 1987. These impulse dummies are entered unrestricted in the VAR. Exclusion of these dummies does not change the results qualitatively except for the failure of the diagnostics (non-normality and/or autocorrelation). Superscripts a, b, and c indicate significance at 1%, 5% and 10%, respectively.

**Table 5: Estimated cointegrating parameters (Johansen and FMOLS methods). Normalised on LYP**

	Johansen, VECM		FMOLS		
Section A: Country-by-country time series results					
	LKP	STR	LKP	STR	$\eta_{\mu}$
GRE	0.431 <sup>a</sup> [0.066]	0.067 <sup>a</sup> [0.023]	0.670 <sup>a</sup> [0.017]	0.040 <sup>a</sup> [0.008]	0.161
IND	0.594 <sup>a</sup> [0.043]	0.054 <sup>a</sup> [0.012]	0.653 <sup>a</sup> [0.031]	0.045 <sup>a</sup> [0.014]	0.065
KOR	0.637 <sup>a</sup> [0.040]	0.168 <sup>a</sup> [0.045]	0.771 <sup>a</sup> [0.015]	0.048 <sup>a</sup> [0.017]	0.105
PHL	0.273 <sup>a</sup> [0.031]	0.007 [0.008]	0.242 <sup>a</sup> [0.042]	-0.007 [0.009]	0.235
SAF	0.487 <sup>a</sup> [0.145]	-0.519 <sup>a</sup> [0.143]	0.612 <sup>a</sup> [0.071]	0.016 [0.062]	0.358
TWN	0.615 <sup>a</sup> [0.035]	0.176 <sup>a</sup> [0.062]	0.720 <sup>a</sup> [0.013]	0.090 <sup>a</sup> [0.023]	0.091
Section B: Between dimension panel results:					
Panel	0.506 <sup>a</sup> [0.046]	-0.008 <sup>b</sup> [0.004]	0.611 <sup>a</sup> [0.021]	0.039 <sup>b</sup> [0.017]	

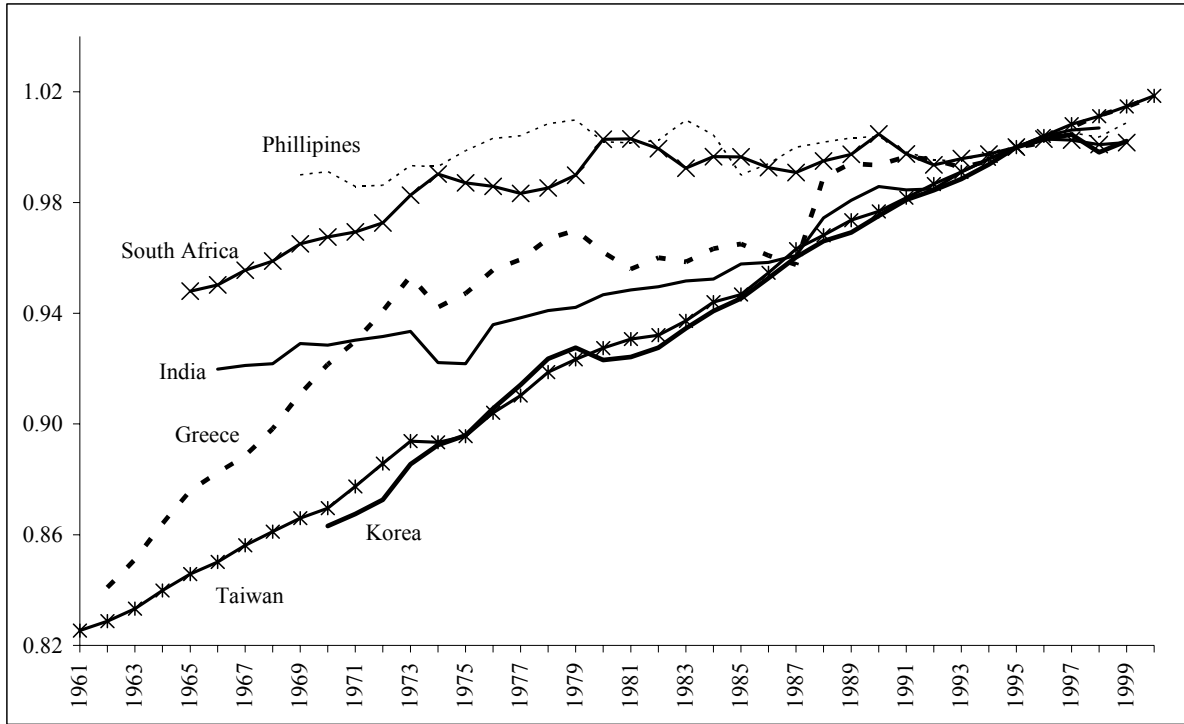
The country mnemonics are: GRE=Greece; IND=India; KOR=Korea; PHL= the Philippines; SAF= South Africa; TWN=Taiwan. Figures within parenthesis [.] are respective standard errors. Bartlet window of second order is used for the FMOLS estimations. The  $\eta_{\mu}$  statistics are the KPSS tests of stationarity of the error correction terms derived from the FMOLS estimates. The null is that the error-correction term is level stationarity. The 5% and 1% critical values for  $\eta_{\mu}$  are 0.463 and 0.739, respectively. Variable definitions are: LYP = log of per capita real GDP; LKP = log of per capita real physical capital stock; and STR = log(CLR/BLR). Superscripts a, b, and c indicate parameter significance at 1%, 5% and 10%, respectively.

Table 6: Tests of equality of panel and country-specific parameters

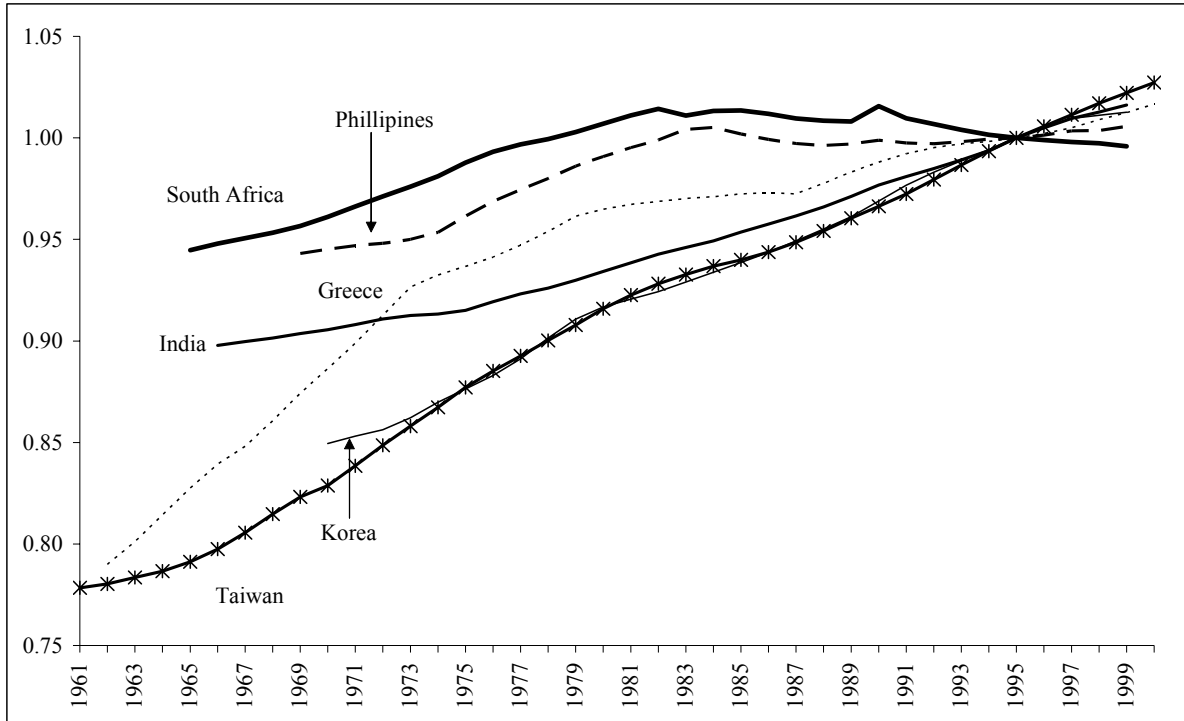
Johansen (LR Statistics)		FMOLS (Wald Statistics)	
LKP	STR	LKP	STR
8.738	30.440 <sup>a</sup>	292.782 <sup>a</sup>	34.149 <sup>a</sup>
DF : (6)	(6)	(6)	(6)

The null is that the country-specific parameters are jointly equal to the panel estimates (parameters). Variable definitions are: LYP = log of per capita real GDP; LKP = log of per capita real physical capital stock; and STR = log(CLR/BLR). DF = Degrees of Freedom which is  $\chi^2(6)$ . Superscript 'a' indicates rejection of the null of equality at 1%.

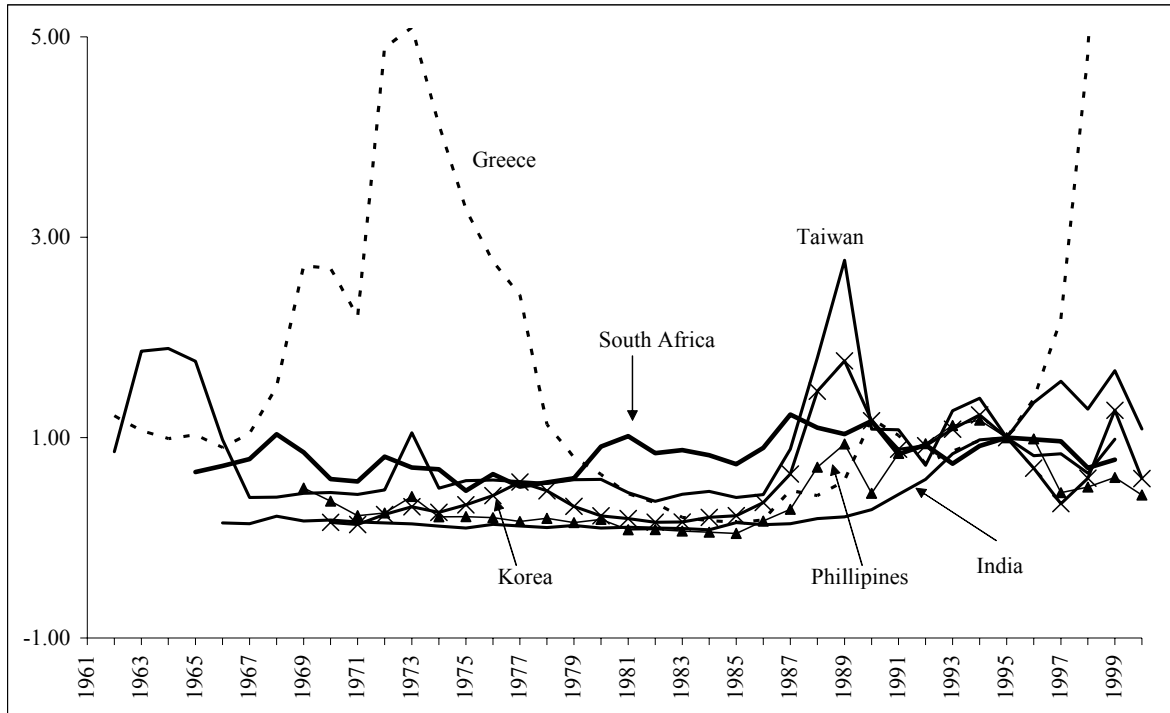
Figure 1: GDP per capita (1995=1)



**Figure 2: Per capita capital stock (1995=1)**

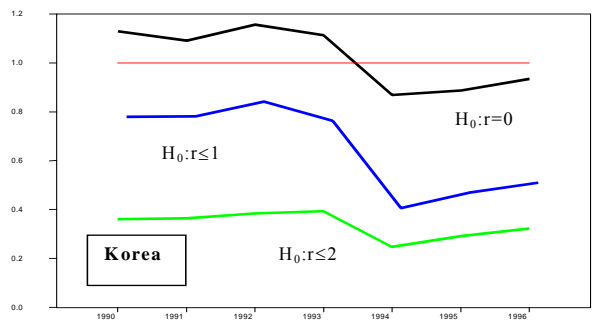
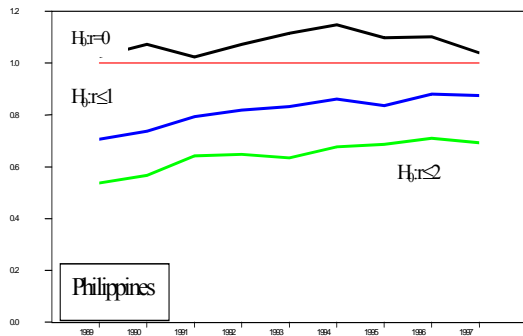
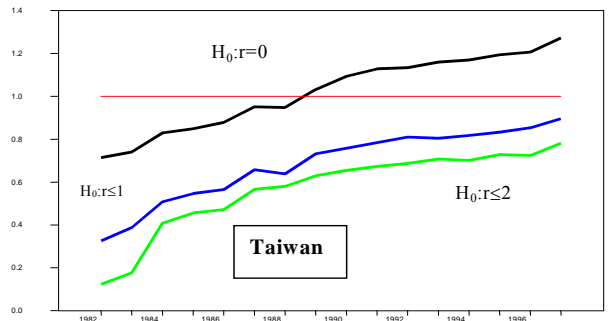
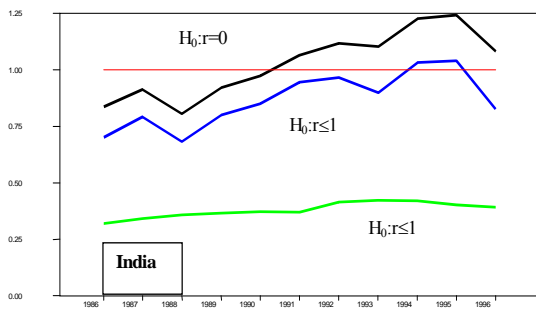
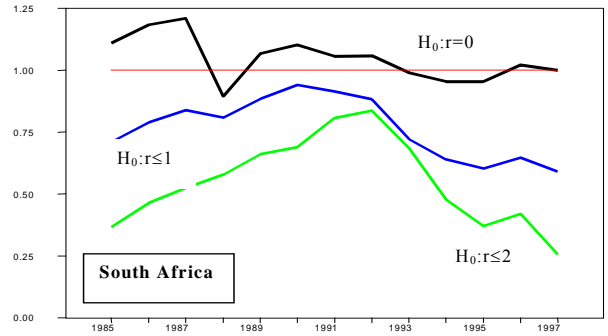
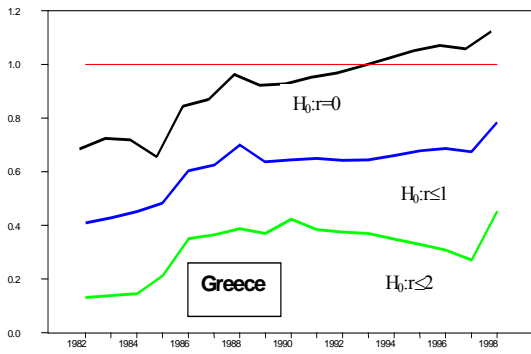


**Figure 3: Financial structure variable (1995=1)**





**Figure 4: Plots of Scaled Recursive LR- Statistics (Rank Stability Tests: R Model)**



**Figure 5: Parameter Stability Tests (Plots of Scaled Recursive LR- Statistics)**

