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IS THERE A LINK BETWEEN PENSION-FUND ASSETS AND ECONOMIC GROWTH? - A CROSS-COUNTRY STUDY

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Abstract: Debate over superiority of pension funding over pay-as-you-go links notably to the question whether funding improves economic performance sufficiently to generate additional resources to meet the needs of an ageing population. To address this issue, we design a modified Cobb-Douglas production function with pension assets as a shift factor, and investigate the direct link between pension assets and economic growth employing a dataset covering up to 38 countries, using a variety of appropriate econometric methods. We find positive results for both OECD countries and Emerging Market Economies (EMEs), with consistent evidence for a larger effect for EMEs than OECD countries.

Key words: Pension funds, economic growth, production function, panel estimation
JEL Classification: G23, O16, C33

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Introduction

The current global demographic shift toward population aging, largely reflecting rising life expectancy and declining fertility (Munnell 2004) has led many countries across the world to re-evaluate their pension systems. Typically, they switch wholly or partially from unfunded systems, e.g. pay-as-you-go (PAYG) to funded systems, with reform in Emerging Market Economies (EMEs) often supported by World Bank finance (Holzmann and Hinz 2005).

Given the funded nature of many new pension schemes, pension fund assets have increased across many countries. Looking first at OECD countries, in 1980, UK pension assets were equivalent to 115.6 billion US dollars (21.5% of GDP), whereas in 2000 these two figures had increased to 1281.5 billion US dollars (79% of GDP) (OECD 2003). The trend was similar in most other OECD countries. Table 1 shows that as of 2000, total pension fund assets across our selected advanced OECD countries were US$12 trillion. The US as the biggest pension market accounted for just above half of the total with Japan and the UK following. In terms of pension assets relative to GDP, the Netherlands had the largest ratio at 149% of GDP, while this figure for New Zealand, at 0.69%, was the smallest across OECD countries.

As regards data for EMEs shown in Table 2, Chile is the country which pioneered reform towards private funded pensions and its experience is often cited to justify funded pension reform; in that country pension funds grew from zero in 1980 to 60 per cent of GDP as of 2002. The biggest EME pension markets were, however, Singapore and Malaysia which adopted publicly managed funded Provident Fund pension systems in the 1950s. Other EMEs with significant pension assets include Brazil and Mexico. Total pension assets across our selected EME countries in 2002 were US$ 280 billion, while the average pension asset-GDP ratio was 12 per cent, much less than that of OECD countries which was 42 per cent.

Given demographic trends and the structure of funded schemes, it is virtually certain that pension funds will continue their rapid expansion during the coming decades. In this context, a key issue in pension reform is whether such a shift from PAYG to funding is largely a matter of reallocation of the financial burden of ageing (with the risk of a generation paying twice), or whether funding improves economic performance sufficiently to generate at least some of the additional resources required to meet the needs of an ageing population. There are several aspects to this question. One is whether funding leads to an increase in saving which permits higher capital formation. A second is whether, independently of the impact on saving, there are favourable effects of funding on the functioning of capital and labour markets, for example via acceleration of financial development, generating in turn a more efficient allocation of capital. A third is whether, following from these effects, a direct impact of funding on growth can be discerned.

Whereas there is quite extensive work on funding’s effect on saving and on financial development (Hu (2005a), Davis and Hu (2005), Davis (2005)), the direct role of pension funds in economic growth has been little examined. Is pension-fund growth positively associated with economic performance? And if so, how long will this positive impact continue? In this paper, we seek to provide insight into these questions with both a theoretical model and related empirical work for most OECD countries and selected EMEs.

The paper is structured as follows. Section 1 provides a brief literature review on the issue of whether and how pension fund growth may impact on economic performance. Section 2 deals with the model specification, which is derived from the Cobb-Douglas production function, and views pension fund assets as a shift factor, an idea developed from McCoskey and Kao (1999) and Arestis et al (2004). Data and variables are discussed in Section 3. In Sections 4, we test our data's stationarity by using unit root tests and find variables to be I(1), implying a need to allow for cointegration. In Section 5, our first econometric work is conducted with the help of cointegrating dynamic OLS (DOLS) model, which we consider most appropriate for the question in hand. Complementing this, in Section 6, we follow the dynamic heterogeneous estimation procedures designed by Pesaran and Smith (1995) to look at the average long run relations and allow for cross country heterogeneity. In Section 7, we move to country-by-country co-integration tests, investigating whether there is a long run relationship between pension funds and economic growth and again allowing again for cross country heterogeneity. Impulse responses of output per capita to pension assets in the related Vector Error Correction Models (VECM) are calculated as well as variance decompositions. In each case we assess results with and without a time trend, which may capture other influences on the production relation such as structural reform.

Summarising the results (see Table 11), a strong and positive relation is found in the cointegrating dynamic OLS panel model between pension assets, output and capital, with larger effects of pension assets on output for EMEs when a time trend is included. Panel cointegration coefficients using mean-group dynamic heterogeneous models again find a positive and significant average long run relationship between pension assets and output, again notably for emerging market economies. Country-by-country cointegration tests typically find a cointegrating relationship between the I(1) variables pension assets/GDP, the capital stock per capita and output per capita. Impulse responses in the related VECM with and without the trend show that a rise in pension assets typically boosts output per worker, and during a 25-year period, the effect typically remains positive. Significantly larger effects are found for EMEs than OECD countries.

1. Literature review

Whereas there is some evidence of a small positive effect of pension-funding on household saving (Kohl and O'Brien 1998), the relevant variable for economic growth is national saving which largely determines investment[1], as would be predicted by a standard Solow (2000) growth model. James (1998) argues that one main advantage of World Bank multi-pillar model of pension reform is that national saving as well as personal saving could be boosted. But any positive effect of pension fund growth on personal saving could be offset at the level of national saving by the impact on public finances of the costs involved in the transition to a privately funded system (Holzmann 1997), as well as the costs of tax subsidies to personal saving. A key aspect of this issue is hence how pension-reforming governments finance existing social security obligations. If the government tries to finance the implicit pension debts by public debts, then public savings would decrease, so the overall national saving rate might be unchanged or even fall.

There is conflicting empirical evidence on this point. Schmidt-Hebbel (1999) estimated that pension reform in Chile raised the national saving rate. Given the difficulty of pinning down how the pension reform was financed in Chile, he considered three cases, i.e. fiscal contraction-based financing of pension reform at the levels of 100%, 75% and 50%. On balance, he suggests that between 10% and 45% of the rise in national saving could be explained by pension reform, with the remaining being explained by structural reform, e.g. tax reform etc. Lopez-Murphy and Musalem (2004) study 50 countries and find that national saving is boosted where pension funds are the result of a mandatory pension programme, but not when they are voluntary. On the other hand, Samwick (1999), working with a panel of

[1] Whereas such relations can be weakened by international capital flows, the extensive literature on the "Feldstein-Horioka puzzle" shows that domestic investment continues to be strongly related to domestic (i.e. national) saving.

countries, found that no countries except Chile experienced an increase in gross national saving rates after pension reform towards non-PAYG systems. He included control variables such as the log of per capita income, per capita income growth, the private credit to income ratio, demographic indicators and the urbanisation rate to avoid omitted variables bias. Furthermore, Bosworth and Burtless (2004) found that OECD countries that seek to prefund social security obligations such as Japan and the US incur offsetting increases in government borrowing that again offset any difference in national saving.

Given the doubt about a link of pension reform to national saving, we consider it important to focus on some alternative channels for pensions to affect growth, via improved economic efficiency and resource allocation. The link between pension funds and capital market development has been widely analysed in the recent literature, as reviewed in Davis and Hu (2005). Both prices and quantities of long term financing may be favourably affected, which may in turn raise productive investment and improve resource allocation.

Focusing on emerging market economies (EMEs), Walker and Lefort (2002) argue that pension funds can decrease the cost of capital via three channels. The first channel is more developed capital market resulting from pension reforms, thus making the issuing of securities cheaper. Secondly, even allowing for short-term performance evaluation (Davis and Steil 2001), the expected investment time horizon of pension funds is longer than that of individuals and firms, thus reducing the ‘term premium’. Third, the equity risk premium is reduced due to pension funds’ pooling and professional management. Both the term premium and risk premium’s reduction might lead to a decrease in the average cost of capital, which spurs investment. In addition, they give some evidence that pension funds reduce security price volatility, implying a lower risk premium for their panel of emerging market economies, although an opposite result is found by Davis (2004) for G-7 countries. Turning from prices to quantities, Catalan et al (2000) give evidence that contractual saving institutions, e.g. pension funds, positively Granger-cause equity market capitalisation as well as value traded, while Impavido et al (2003) and Hu (2005a) find a positive relationship between contractual saving assets and bond market capitalisation/GDP.

In sum, the current literature suggests a positive relation between pension fund growth and financial development, see also the survey in Davis (2005). Given it is widely considered that financial development is positively associated with economic growth (Levine and Zervos 1998; Beck and Levine 2004), then pension funds might enhance economic growth via their impact on financial development, independently of an effect on national saving.

Pension funds may also boost economic growth via improved corporate governance (Clark and Hebb 2003; Myners 2001)[2]. Clark and Hebb (2003) identify four drivers which facilitate pension funds’ corporate engagement, which they see as foreshadowing the so-called “Fifth Stage of Capitalism”. The first driver is the widespread use of indexation techniques in the pension funds industry, which hinders “exit” via sale of shares in underperforming companies which are in the index. The second driver is the increasing demand by owners for more transparency and accountability, particularly after the Enron, Worldcom and Parmalat scandals. Third, there is pension funds’ pressure to undertake socially responsible investing (SRI). Fourth, pressures to “humanize” capital with social, moral and political objectives extend pension funds’ simple concerns for rate of return.

A positive impact of pension fund activism on corporate performance at the firm level is well

[2] The effectiveness of pension funds’ positive impact on corporate governance has been challenged by Orszag (2002) and empirical works in the US such as Del Guercio and Hawkins (1999).

documented, although empirical work is largely focused on the US[3]. But our concern in this paper is whether pension fund growth is a potential engine of economic growth via its effect on corporate performance at the macro level, an issue which is ignored or dismissed by most current pensions research. An exception is Davis (2002, 2004) who argues that complementary studies at the macro level are needed, because effects of governance initiatives from institutional investors may go wider than the "target firms" to the whole economy. This is because unaffected firms have natural incentives to improve their performance so as to avoid the threat from pension fund activism in the future (Marsh 1990). Therefore, if a significant proportion of firms, whether directly affected and indirectly affected by pension fund activism, tend to improve their performance, the overall effect might be higher economic growth and productivity for the whole economy. Consistent with this, Davis found inter alia that institutional holding of equity is related to a boost to Total Factor Productivity (TFP) and in R and D.

Besides the issue of corporate governance, labour market performance is relevant. It is well known that due to the weak link between pension contributions and benefits under defined-benefit PAYG systems, there is a tendency towards earlier retirement and job immobility. For example, over the postwar period, there was a very sharp fall in participation rates for those men over state pension age (65+) in EU countries (Disney 2002). One contributing factor was the disincentives imbedded in public pension systems (Blondal and Scarpetta 1998). In addition to the pension system's impact on labour supply, Disney (2003) argues that the distortionary "tax component" of public pension contributions can also affect labour demand if the employee can pass through the burden of pension contribution to consumers for example via product prices, because product demand falls and producers might consider reducing the demand for labour. In view of such problems, James (1998) suggests that "the close linkage between benefits and contributions, in a defined-contribution plan is designed to reduce labour market distortions." In consequence, economic growth might be increased, e.g. due to a higher labour participation rate after pension reform. Such effects might be smaller where defined benefit funded schemes predominate.

Looking at the direct link of growth to pension reform, most extant studies have focused on Chile. Holzmann (1997) found a positive relationship between pension reform and economic growth. With the simple Solow residual specification of TFP, it was found that improving financial market conditions following the pension reform significantly positively affected TFP. But this model suffers from low "t" values which might result from multicollinearity between independent variables, e.g. the unemployment rate and stock market index. Meanwhile, Schmidt-Hebbel (1999) reached the conclusion that pension reform in Chile boosted private investment, the average productivity of capital and TFP. One single regression was estimated to obtain the coefficients of parameters, then these coefficients are used to calculate the rise of each variable attributed respectively to structural reform, (e.g. tax reform) and pension reform. In all, he concluded that pension reform in Chile had a positive impact on the private investment rate, average productivity of capital and the TFP growth rate. For example, pension reform contributed to 0.1-0.4 per cent of the 1.5 per cent increase in TFP growth rate, while 0.4–1.5 per cent of the total 13 per cent rise in private investment rate was attributed to pension reforms with the remainder being explained by structural reform.

Empirical work which investigates the direct link between pension fund growth and economic growth at a transnational level is quite scarce to our knowledge, although Davis (2002 and 2004) with a dataset covering both pension funds and life insurance companies, looked at the relation between institutionalisation and economic performance at the macro level. Although

[3] See Wahal (1996), Smith (1996), and more recent work by Woikdtke (2002) and Coronado et al (2003) for estimates of the impact of pension activism on corporate performance at the firm level.

his results are, as noted, consistent with higher TFP, he finds no direct effect of the proportion of equity held by life insurers and pension funds on GDP growth. Again, Davis (2004) using a dataset of 16 OECD countries and a standard Levine-Zervos (1998) specification for finance and growth does not find a positive direct link between institutionalisation (life insurance and pension assets/GDP) and growth per se.

On the other hand, using the technique developed by Hurlin and Venet (2003) and Hurlin (2005), Hu (2005b) shows that Panel Granger Causality tests do indicate homogeneous causality from pension assets to GDP growth in 38 countries as well as in the subgroups OECD (18 countries) and EMEs (19 countries). Reverse causality is weaker, and notably for emerging markets there is no strong evidence that GDP growth homogenously causes pension assets.

Taking into account the above literature review, this paper seeks to contribute to the current growth and pensions literature in three areas. First, we design a modified Cobb-Douglas production function with the inclusion of pension assets as a shift factor. Second, we employ a set of different econometric methods to test the model on data for up to 38 countries, which includes a cointegrating dynamic OLS estimator for the main panel results and also the dynamic heterogeneous models designed by Pesaran and Smith (1995) to look at the average long run relations between variables, allowing for cross country heterogeneity. Third we directly link pension assets to economic growth in a co-integration relationship on a country-by-country basis and investigate the extent to which they are correlated in the long run as well as the impulse responses and variance decomposition in the related Vector-Error-Correction Model.

2. Model specification

The Cobb-Douglas production function is widely used in the economic literature:

$$Q = AK^{\beta}L^{1-\beta} \tag{1}$$

where A is technology, K is the capital stock and L is the labour force. Generally, the Cobb-Douglas function is specified as shown in Equation (1). But in this study, we modify the function slightly so as to facilitate our analysis of the implication of pension fund assets for output Q. In addition, in view of our panel analysis, we use a double subscript on its variables.

$$Q_{i,t} = A_{i,t} \times (P_{i,t})^{\lambda_i} \times (K^{\beta_i}{}_{i,t}) \times (L^{1-\beta_i}{}_{i,t}) \tag{2}$$

where: i: time series dimension;
t: cross section dimension;
Q: aggregate output, proxied by GDP;
A: state of technology;
P: pension funds, proxied by pension fund assets/GDP;
K: capital stock[4];
L: labour supply, proxied by total population;
λ: elasticity of aggregate output with respect to pension fund assets;
β: elasticity of aggregate output with respect to the capital stock.

Equation (2) suggests that aggregate output is affected both by technology A and pension fund assets P, which act as shift factors, as well as capital K and labour L. Note that the model does

[4] Capital stock is calculated based on the perpetual inventory method. Consistent with Luintel and Khan (1999), we used 8 per cent of depreciation rate and averaged first 3-year growth rate to obtain the initial capital stock.

not assume pension fund growth raises saving – trends in national saving will be captured by the capital stock variable, to the extent external balance is maintained. In effect, we test whether owing to better resource allocation, incentives etc., pension fund growth makes the capital stock more productive. Arestis et al (2004) and McCoskey and Kao (1999), among others, use the similar specification, i.e. a generalised Cobb-Douglas production function with relevant additional variables such as urbanization rates or the nature of the financial system set as shift factors into the standard function. Technology may then be specified as follows:

$$A_{i,t} = e^{\alpha_i + \gamma_i t + \varepsilon_{i,t}} \qquad (3)$$

This specification is in line with McCoskey and Kao (1999), where α is the intercept, t is the time trend and ε is the residual term. Specifying the state of technology in this way assigns each of our country sample with the country-specific intercept and time trend (allowing for heterogeneity across countries that might relate to factors such as structural reform) and also introduces a stochastic element, i.e. ε into the model as indicated in Equation (5) below. Replacing technology A in Equation (2) by its expression in terms of t as shown in Equation (3) gives

$$Q_{i,t} = e^{\alpha_i + \gamma_i t + \varepsilon_{i,t}} \times (P_{i,t})^{\lambda_i} \times (K^{\beta_i}{}_{i,t}) \times L^{1-\beta_i}{}_{i,t} \qquad (4)$$

Then, normalising by $L_{i,t}$ and taking logs from both sides in Equation 4, we have

$$\frac{Q_{i,t}}{L_{i,t}} = e^{\alpha_i + \gamma_i t + \varepsilon_{i,t}} \times (P_{i,t})^{\lambda_i} \times (\frac{K_{i,t}}{L_{i,t}})^{\beta_i}$$

$$Q^*{}_{i,t} = e^{\alpha_i + \gamma_i t + \varepsilon_{i,t}} \times (P_{i,t})^{\lambda_i} \times (K^*{}_{i,t})^{\beta_i}$$

$$LnQ^*{}_{i,t} = \alpha_i + \gamma_i t + \lambda_i LnP_{i,t} + \beta_i LnK^*{}_{i,t} + \varepsilon_{i,t} \qquad (5)$$

$$\text{where } Q^*{}_{i,t} = \frac{Q_{i,t}}{L_{i,t}} \quad and \quad K^*{}_{i,t} = \frac{K_{i,t}}{L_{i,t}}$$

$$\gamma_i = \lambda + \omega_{1i} \quad , \quad \lambda_i = \lambda + \omega_{2i} \quad and \quad \beta_i = \phi + \omega_{3i}$$

$Q^*{}_{i,t}$ is output per worker and $K^*{}_{i,t}$ is capital per worker. The model shown in Equation (5) is the standard formulation of Swamy's Random Coefficient Model (RCM) (Swamy and Tavlas 1995) where we can allow for heterogeneity across countries in terms of time (t), pension fund assets (LnP) and capital per worker (LnK). We view this model as appropriate in that pension fund assets' impact on output might show marked differentials across countries.

Following the model above, we regress capital per worker (CPW) and pension fund assets/GDP (PFAGDP), which are K* and P in Equation 5 respectively, on output per worker (OPW) or Q*, using various econometric techniques. We estimate with and without the time trend (t) which may capture other influences on production relations such as structural reforms.

Following Arestis et al (2004) we do not include some of the standard variables typically entered in cross-sectional cross country growth regressions such as years of schooling, as well as corruption, social capital, inequality and rule of law. On the one hand, it would not have been feasible to build an annual time series for these variables. Furthermore, we consider that a generalized production function estimated is the appropriate specification for the issue in hand and using panel data with fixed effects and a time trend (in some specifications) will capture any relevant differences in growth performance across countries.

3. Data and variables

Before describing estimation, we outline issues in data construction and unit root tests. Regarding the calculations of Q* and K* we use standard macro-economic data from the World Development Indicators 2003 (WDI) database. The capital stock is derived by the perpetual inventory method. Consistent with Luintel and Khan (1999), we used an 8 per cent depreciation rate and averaged the first 3-year growth rate to obtain the initial capital stock.

Pension fund asset data were collected from a variety of sources. For OECD countries, OECD (2003) and Davis and Steil (2001) are the main sources, but some are expanded and updated by checking financial statistical reports in individual countries, e.g. National Financial Statistics for the UK data and Institute of Pension Research and Nikko Financial Intelligence, Inc for the Japan data. For Latin American countries, the website of Federación Internacional de Administradoras de Fondos de Pensiones (FIAP) (International Federation of Pension Fund Administrations) in Chile provides pension data up to the year end of 2003 on many Latin American countries. For South Asian countries and South Africa, pension data are largely compiled individually by searching local central banks' Financial Bulletins, although ASEAN Social Security Association's website was used to update recent pension data on some Southeast Asian countries.

Regarding the data observation period, in general, for capital per worker and output per worker we have data for years between 1960 and 2002. But pension data are an exception. For OECD countries, e.g. the UK, the US, we have data ranging from 1960s to 2002, while for many EMEs, e.g. Brazil, the data available are relatively limited. See Appendix 1 for details of the variables across our 38 countries.

4. Panel unit root test

Before proceeding to formal panel regression analysis, the first step is to examine our data's stationarity.

4.1 Specification of tests

There are a number of ways to test panel data's stationarity (Maddala and Wu 1999; Baltagi 2001). In this study, in order to check our results' robustness, we use three different but commonly quoted tests, i.e. one designed by Levin, Lin and Chu (2002) (hereafter LLC), one by Im, Pesaran and Shin (2003) (hereafter IPS), and last one by Hadri (2000).

Consider the following model

$$y_{i,t} = \rho_i y_{i,t-1} + X_{i,t}\delta_i + \varepsilon_{i,t} \quad i = 1,...N : t = 1,...T \tag{6}$$

where y is our variable of interest; X is a vector of exogenous variables, including fixed effects and/or a time trend, or simply a constant, based on the modelers' assumptions. $\varepsilon_{i,t}$ are i.i.d. $(0,\sigma_\varepsilon^2)$. As customary, t proxies time, while i proxies country.

The principal difference between LLC and IPS is the assumption made on ρ_i. LLC proposes that $\rho_i = \rho$, implying the coefficient of lagged dependent variable in Equation (6) is the same across countries, while under IPS, ρ_i is allowed to vary across countries. Given that in our sample, both OECD countries and EMEs are included, we put more emphasis on the latter test, i.e. IPS (2003), in that there might be heterogeneity across countries.

Both LLC and IPS tests are an extended version of time series' Augmented Dickey-Fuller test (ADF) into the context of panel data. The formulation is as follows:

$$\Delta y_{i,t} = \beta y_{i,t-1} + \sum_{j=1}^{p_i} \rho_{i,j} \Delta y_{i,t-j} + X_{i,t}\delta_i + \varepsilon_{i,t} \quad i = 1,...N : t = 1,...T \tag{7}$$

LLC tests the null hypothesis of β =0, while IPS is testing that of β_i =0 for all i. In addition, for the IPS test, t-bar statistics is used, which are formed as a simple average of the individual t statistics for testing β_i =0 in Equation 7, namely

$$t - bar_{NT} = N^{-1} \sum_{i=1}^{N} t_{iT} \tag{8}$$

Both LLC and IPS are commonly used in the current empirical literature for panel data. It has been argued, however, that they both suffer from the lack of power (Hadri 2000). In other words, the null hypothesis of a unit root tends to be accepted or not rejected unless there is strong evidence to the alternative, one form of type II error (Davidson and MacKinnon 1993; Greene 2003). Therefore, it is suggested to test a null of stationarity as well as a null of a unit root. One well-known test for the null of no unit root is that proposed by Hadri (2000). Hadri testing is a residual based Lagrange multiplier (LM) test. Consider the model,

$$y_{i,t} = r_{i,t} + \beta_i t + \varepsilon_{i,t} \tag{9}$$

where $r_{i,t} = r_{i,t-1} + \mu_{i,t}$, a random walk. The LM statistic is formulated as follows:

$$LM = \frac{\frac{1}{N}\sum_{i}^{N} \frac{1}{T^2} \sum_{t=1}^{T} S_{i,t}^2}{\hat{\sigma}_\varepsilon^2} \tag{10}$$

where $S_{i,t} = \sum_{j=1}^{t} \hat{\varepsilon}_{i,j}$ and $\hat{\sigma}_\varepsilon^2 = \frac{1}{NT} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{\varepsilon}_{i,t}^2$

$\hat{\varepsilon}_{i,j}$ is the estimated residual from Equation (9), $S_{i,t}$ is the partial sum of residuals, while $\hat{\sigma}_\varepsilon^2$ is the estimate of the error variance. Hadri's residual-based LM test for the null of stationarity is promising in that it increases the power of testing for the null of a unit root. One problem associated with Hadri (2000), however, that like LLC (2003), it assumes the homogeneity of coefficients of $\rho_i = \rho$ in Equation (6). As we mentioned earlier, in our study, we use a dataset covering both OECD countries and EMEs; therefore, an assumption that ρ varies across sections might be more appropriate.

4.2 Results for panel unit root tests

Table 3 presents the results of the panel unit root tests. For the log of the pension assets to GDP ratio (PFAGDP) our results, under all three testing approaches, are in favour of non-stationarity in levels, and stationarity in first differences, implying that PFAGDP is an I(1) variable. Regarding the log-levels of output per worker (OPW) and capital per worker (CPW), under IPS and LLC, the null hypothesis of non-stationarity could not be rejected for this panel of 38 countries. But after first differencing, the null hypothesis of non-stationarity is rejected and the alternative hypothesis of stationarity be accepted. This is consistent with our assumption that OPW and CPW are also I(1) series.

By employing the Hadri (2000) test, however, we could reject the hypothesis of no unit root under both levels and first differences. After second differencing, OPW and CPW become stationary, as the null of stationarity could not be rejected. This is intriguing and implies that OPW and CPW are I(2) variables if only based on Hadri. But, it is worth noting again that Hadri (2000) assumes a common unit root process, which as we have motioned earlier is less relevant in this study. Therefore, together with other two testing procedures, we believe PFAGDP, CPW and OPW are all non-stationary and I(1) variables.

5. Dynamic OLS estimation

5.1 Econometric specification

In this section, we seek to identify the relation between pension assets and output in the context of our theoretical model by using the dynamic OLS (DOLS) cointegrating panel estimator. In panel data, Kao (1999) finds that the ordinary least squares (OLS) estimator is biased, in that the t-statistics diverge so the inference is not reliable. The fully modified OLS (FMOLS) estimator is argued to be able to correct such bias in certain cases. The FMOLS was first proposed by Philips and Hansen (1990), and extended to the context of heterogeneous panels by Pedroni (1997), and then developed further in Philips and Moon (1999). Based on the simulation results from the Monte Carlo experiments, Kao and Chiang (2000), however, prove that under both contexts of homogeneous and heterogeneous panels, dynamic OLS (DOLS) is superior to fully modified OLS (FMOLS) and other OLS estimators. The advantages of DOLS over FMOLS are no requirement for initial estimation and non-parametric correction[5]. The DOLS model, used in our paper and following Stock and Watson (1993) is as follows:

$$Y_{i,t} = \alpha + \beta X_{i,t} + \sum_{j=-n}^{n} \gamma \Delta X_{i,t} + \varepsilon_{i,t} \qquad (11)$$

where i and t are country and time indices as conventional. $Y_{i,t}$ is the dependent variable, i.e. log output per worker (OPW). $X_{i,t}$ is a vector of explanatory variables, i.e. log pension fund assets/GDP, and log capital per worker (CPW). $\Delta X_{i,t}$ is the first difference of $X_{i,t}$, thereby introducing dynamic structure into the equation. The coefficients of $X_{i,t}$ give the accumulative/total effects. In addition, the length of leads and lags for $\Delta X_{i,t}$ has to be defined. The inclusion of these nuisance parameters in Equation 11 means we can obtain coefficient estimates with satisfactory limiting distribution properties (Kao and Chiang 2000; Kao et al 1999). As mentioned by Kao and Chiang, however, it is difficult to choose the optimal length of leads and lags, which is a major drawback of the DOLS estimator. But, the practice is to use 1 and/or 2 leads and lags.

[5] We also consider DOLS to be more appropriate than system Generalised Method of Moments (GMM) of Arellano and Bover (1995), since GMM is most appropriate when N is large and T is small (Bond 2002). But in our dataset, neither is the case; for example, we only have data covering 38 countries, while observations range from 5 years to 35 years.

5.2 Empirical results

Results are given in Tables 4a and 4b, where our main focus is on the sign, size and significance of the variable LPFAGDP, the log of the pension fund assets/GDP ratio, which indicates the shift in the production function. As noted above, it is arbitrary to choose the length of leads and lags in the DOLS model, but the practice is to use 1 or 2 leads/lags (Mark and Sul 2002, and Kao et al 1999). In this paper, in order to check the robustness of DOLS model as in Pelgrin et al (2002), we used both 1 lead/lag and 2 leads/lags. We split our dataset according to two dimensions, i.e. OECD/EMEs, and with trend/no trend.

Use of a trend is consistent with McCoskey and Kao (1999) where they use a time trend to identify the potential beneficial effect of technological advances on growth over time, as well as structural reform not related to pensions. In addition, as we have noted earlier, the variable capital per worker (CPW) is not stationary even after first differencing based on Hadri test, which might be due to the presence of a deterministic trend. Therefore, the specification with a trend utilised here might be able to deal with this issue. In order to allow for our data to have a deterministic trend as well as to allow for the potential effect of technological advances, we specified a model with a trend as well as without both in this section and for the Mean-Group and Johansen results reported below..

As regards the coefficient of LPFAGDP, in Table 4a where we used 1 lead/lag of the dynamic terms, five out of six estimates are significant and positive as expected, covering all three country groups. In Table 4b where we used 2 leads/lags, results are similar. In each case, for the All-countries estimation, the estimate without the trend is positive and statistically significant, while the estimate with the trend is insignificant, suggesting heterogeneity, which is manifest when the time trend is included. All the EME and OECD estimates are positive and significant. Meanwhile, the time trend term tends to be positive and significant under all cases. It implies that technological advances and structural reforms over time improve the relation of capital and labour to output. Its inclusion means the pension variable is not proxying an omitted trend. The estimates for LCPW are very satisfactory, in that all are statistically significant at 1 per cent, and positive at the range of 0.3-0.8. Finally, the adjusted R-square ratios are quite high in all cases. Note that differences in the size of the coefficients between the 1 and 2 lead/lag specification may relate largely to the difference in country composition, where the former uses data from 37 countries, the latter from 33.

As regards the size of the LPFAGDP coefficients, they are in each case smaller when the time trend is included, implying that there are technical and structural changes that the trend is capturing, which is otherwise incorporated in the pension assets variable. But as noted, for OECD and EME groups and for each lag specification, the coefficients are significant and positive with the trend as well as without it. This parameter measures the total or cumulative effect of pension assets on output. Therefore, it implies that a one percent increase in LPFAGDP raises LOPW by a minimum 0.012 per cent under the case of OECD-with trend, and a maximum 0.068 per cent under the OECD-no trend as in Table 4b. Comparing the OECD and EME results, the OECD pension variable tends to be larger when the trend is omitted but smaller than in EMEs when the trend is included, where the latter results are more plausible. One would expect larger coefficients for EMEs, as is generally the case throughout our results, including the Mean-Group and Johansen regressions reported below. Such a finding is consistent with economic convergence theory (Sala-I-Martin 1996), i.e. poor countries are expected to grow faster than rich countries, as well as recent empirical results by Beck and Levine (2004) and Beck et al (2000) implying financial development is more beneficial to economic growth in EMEs.

The basic results are estimated by unbalanced-panel GLS with fixed effects and cross section weights. To check robustness, we sought to re-estimate with the Seemingly-Unrelated Regression (SUR) technique, which allows for correlations in the error terms. This was not feasible for the All or the EME group, because many of the observation series were too short. However, as shown in Tables 4a and 4b, it is apparent that for the OECD group, using SUR does not change the parameter estimates for LPFAGDP markedly, implying that our result of a clear "shift effect" in the production function from pension funding is a robust one.

6. Dynamic heterogeneous models

In view of the possibility that the impact of pension funds on economic growth may vary across countries, and also consistent with the suggestion of McCoskey and Kao (1999), we in this section seek to look further at the long run relationship by employing dynamic heterogeneous models. Pesaran and Smith (1995) present a number of different estimation procedures for estimating a dynamic panel data model across heterogeneous countries, namely the mean group estimator, aggregate time-series estimator, pooled mean group estimator and cross-section estimator. Due to other approaches' limitations[6] as well as data availability, we use only the mean group estimator in this section, investigating the average long run coefficients.

6.1 Mean group estimator specification

The dynamic model we use in this section is specified as follows:

$$LnQ^{*}{}_{i,t} = \alpha_i + \gamma_i t + \varphi_{1i} LnQ^{*}{}_{i,t-1} + \lambda_{1i} LnP_{it} + \beta_{1i} LnK_{i,t}{}^{*} + \varepsilon_{i,t} \tag{12}$$

Equation 12 is the standard formulation of a dynamic heterogeneous panel model, consistent with Pesaran and Smith's (1995) specification. However, with the consideration of saving degree of freedom, we include only one lag of the dependent variable on the right hand side of the function rather than adding lag one of all independent variables like the autoregressive distributed lag (ARDL) estimation used by Pesaran and Smith (1995). Pesaran (1997) and Pesaran and Smith (1999) argue that the use of the ARDL estimation procedure has advantages over the fully-modified (FM) OLS estimator designed by Philips and Hansen (PH) (1990) for time series co-integration relations, e.g. in that the tests based on PH method have a clear tendency to over-reject in small samples and also show larger bias.

Based on the mean group estimation procedure, we ran regressions for each individual country, then averaged across countries using two methods to obtain the average long run coefficients. According to the first method, the long-run elasticities of LnQ* with respect to LnP and LnK* can be calculated using the formula, $\eta_i = \hat{\lambda}_i \big/ (1-\hat{\varphi}_i)$ and $\xi_i = \hat{\beta}_i \big/ (1-\hat{\varphi}_i)$ respectively. $\hat{\lambda}_i$, $\hat{\varphi}_i$ and $\hat{\beta}_i$ are the estimated values of the corresponding parameters in Equation (12). Then the average long-run coefficients in terms of LnP and LnK* can be computed as

[6] For example, the pooled estimator assumes that the coefficients are homogeneous across sections, an assumption which we wish to ease here.

$\eta = N^{-1}\sum_{i=1}^{N}\eta_i$ and $\xi = N^{-1}\sum_{i=1}^{N}\xi_i$ respectively.

The second method, as presented by Pesaran and Smith (1995), maintains that the average long-run coefficients can also be calculated using the means of short-term coefficients, namely

$$\eta = \overline{\lambda} \big/ 1-\overline{\varphi} \quad \text{and} \quad \xi = \overline{\beta} \big/ 1-\overline{\varphi}$$

where

$$\overline{\varphi} = N^{-1}\sum_{i=1}^{N}\hat{\varphi}_i\,, \quad \overline{\lambda} = N^{-1}\sum_{i=1}^{N}\hat{\lambda}_i \quad \text{and} \quad \overline{\beta} = N^{-1}\sum_{i=1}^{N}\hat{\beta}_i$$

The significance levels or t-values of η_i and ξ_i were calculated by following the formulas, $t-value_{\eta} = \hat{\eta}_i \big/ se(\hat{\eta}_i)$ and $t-value_{\xi} = \hat{\xi}_i \big/ se(\hat{\xi}_{i\,i})$ respectively, where the standard errors were computed as the square root of the variance of $\hat{\eta}_i$ and $\hat{\xi}_i$ divided by the number of groups (Smith and Fuertes 2004).

6.2 Empirical result

Results for individual country coefficients with a time trend are given in Appendix 2, where we ran the regression in Equation 12 on 16 countries[7] individually, i.e. 11 advanced OECD countries and 5 EMEs. The coefficients of LCPW and LPFAGDP measure the short-run effects on output, while these coefficients divided by one minus lag one of output LOPW(-1) measure the long run effects on output. Not surprisingly, results vary across countries. The general pattern, however, is clear. The impact of the capital per worker ratio is generally positive, in 15 out of 16 estimates, indicating the positive impact of capital accumulation on output, and is significant in 7 cases. Regarding the pension assets/GDP ratio, 11 out of 16 estimates show a positive sign, although some are insignificant. Meanwhile, the long run effect of LCPW (the ratio of the coefficient on LCPW to one minus the coefficient on LOPW(-1)) is generally around 1, and the pension asset variable is usually around 0.1. The average short-run coefficients for all explanatory variables are given in the bottom-right corner of Appendix 2. A one per cent increase in pension assets leads to an immediate rise in output by 0.022 per cent, while capital’s contribution is larger at 0.283 per cent. The average lagged dependent variable is 0.718.

Further justified in our approach by the differentials across countries as revealed in Appendix 2, we followed the approach noted above by Pesaran and Smith (1995) to assess the long run relation between output, pension assets and capital. Results, according to the mean-group estimators using Methods 1 and 2 are summarized in Tables 5a, 5b and 5c. As in Section 5, we ran three separate regressions by country groupings, i.e. all 16 countries, 11 OECD countries and 5 EMEs. Table 5a presents results for the ARDL with time trend, based on all 16 countries, while Table 5b is based on 10 countries, excluding Canada, Japan, Malaysia, South Africa, Sweden and Switzerland. We dropped those countries since most of coefficient estimates (at least 3 out of 4 estimates) for those countries are not significant (See Appendix 2 for details). Therefore, their presence might distort our results from the mean-group estimators. One of the

[7] 22 other countries were excluded due to the small number of observations.

reason pension assets ratios are insignificant in those countries might be the simple ARDL model we specified. However, in order to keep the specification consistent across countries, and to follow the methodology by Pesaran and Smith (1995), we retain it in this section. Finally in Table 5c we show corresponding results for all 16 countries without the time trend.

Results in Table 5a are satisfactory and encouraging, as all estimates under the two methods and three groups are positive, indicating a positive average long run relationship between pension assets, capital stock and output. For example, for OECD countries, a one per cent increase in the capital stock raises output by 0.936 per cent under method 1 and 0.947 per cent under method 2. These two estimates are quite close to each other. In fact, it is this estimation robustness that leads us to use the simplified model compared with Pesaran and Smith (1995).

Concerning the log of pension assets/GDP, we find that All countries and EMEs have highly significant coefficients, with the long run effect being around three times larger in EMEs than in the All country average. Note however, that this is strongly affected by the result for Chile, without which the EME result would be similar to that from DOLS set out in Section 5.2. Whereas the estimates for OECD countries under both methods in Table 5a are not significant, as noted above, some country by country results feature largely insignificant coefficients. In order to address this problem, we excluded those countries, and the subsequent results are presented in Table 5b. We still have the expected signs and all the LPFAGDP variables are now significant and positive. The effect is, unsurprisingly, larger for EMEs than OECD countries as well as than All countries.

The third set of results in Table 5c are for the equations without the trend. Here we find that for all 16 countries, there is a significant and positive effect of LPFAGDP, thus supporting the result with trend. The coefficients are larger than with the time trend for OECD countries, but reflecting the result for Chile, they are smaller for All countries and EMEs. The EME coefficients are again consistently larger than for OECD countries.

7. Co integration test

As noted in Section 4, pension fund assets, capital per worker and output per worker are all I(1) variables, then we may be interested in whether they are co-integrated, i.e. whether there exists a long run relationship between them. We address this issue in this section on a country-by-country basis to allow for heterogeneity, as well as calculating panel estimates of the cointegrating coefficients.

A co-integrating relationship captures the long run or equilibrium relationship between non-stationary, i.e. I(1) variables. If variables are non-stationary, particularly in the case of time series data, but the common residual terms are stationary, i.e. I(0), then we say these variables are co-integrated and economic theory as set out in Section 2 suggests forces which tend to keep such series together, and do not let them drift too far apart (Banerjee et al 1993). In addition, if variables are co-integrated, our estimates are super-consistent. In other words, our estimates are not only consistent, but also converge to their true values more quickly than normal (Davidson and MacKinnon 1993).

7.1 Specification

In this paper we employ the VAR-based cointegration test (Vector Error Correction Model) using the methodology developed by Johansen (1991 and 1995). The specification is as follows:

$$y_t = A(L)y_{t-1} + \varepsilon_t \qquad (13)$$

where $A(L) = A_1 + A_2 + ... + A_k L^{k-1}$

y_t is a k-vector of I(1) variables, i.e. OPW, CPW and PFAGDP in this paper. L is the lag operator, and the lag order is selected based on a range of information criteria, i.e. AIC (Akaike information criterion) and SC (Schwarz information criterion). Generally, the suggested lag order is 2 years, although in some cases it extended to 3 years. If Equation (13) is written as VAR format, then we have

$$\Delta y_t = \Gamma(L)\Delta y_{t-1} + \Pi y_{t-k} + \varepsilon_t \qquad (14)$$

where

$$\Gamma_i = -(1 - A_1 - ... - A_i), i = 1,...k-1$$
$$\Pi = -(1 - A_1 - ... - A_i) \text{ or } \Pi = \alpha * \beta'$$

where α is the speed of adjustment from short run deviation to long run equilibrium, and β' is the cointegrating vector, which thus represents the long run coefficients. Based on Granger's representation theorem, the Johansen VAR-based cointegration test is to first estimate the Π matrix from an unrestricted VAR and then test whether the restriction suggested by the reduced rank of Π - the number of cointegrating relations - is rejected.

7.2 Results for Cointegration test

This section presents the estimation results for the Johansen cointegration test. Again, we consider two slightly different specifications, i.e. one without a trend and the other with trend. We group our sample into OECD countries and EMEs, which in turn are estimated separately.

Tables 6a and 6b give results of our first specification, i.e. without a trend. In most cases, the Trace and Maximum-Eigenvalue statistics indicate a co-integrating relationship between our variables, and the null hypothesis of non-cointegration is rejected at either the 5% or 10% level. Note that the signs of coefficients are opposite to those of the impact of the variable on LOPW because the equations are normalized in the form 0 = LOPW - a_1LPFAGDP - a_2LCPW

As shown in Table 6a, in only two of eleven OECD countries, i.e. Canada and Switzerland is the sign of coefficients on LPFAGDP positive, implying a negative relationship between pension assets growth and economic output in the normalized cointegrating relation. For all the other countries, however, the sign is negative, as expected. For almost all of these countries, pension fund growth has a statistically significant and positive relationship with output per worker, the extent of which varies from 0.04 for Sweden to 0.27 for Germany. The small size of the positive effect in Sweden could also be due to the restriction of Swedish's ATP scheme from equity investment and state management of the fund (Davis 2003).

Regarding the other regressor, i.e. LCPW (capital per worker), our estimates are satisfactory, as all coefficients are negative, implying a positive linkage between economic output and the capital stock across OECD countries. In addition, the estimates of coefficients of LCPW are quite close to each other; for seven out of eleven countries, it is between 0.55-0.80, implying a comparable production function relationship among developed OECD countries. All estimates except in Canada, Sweden and Switzerland, are less than 1, consistent with our model in

Section 2, which suggests that the β-elasticity of aggregate output with respect to capital should not be greater than 1.

Results for EMEs are given in Table 6b. All coefficient estimates for LPFAGDP except for Malaysia are negative. Therefore, a beneficial effect of pensions on growth is also found across EMEs. For example, for Chile, one per cent increase in pension assets can contribute to economic growth by 0.14 per cent; this complements findings by Schmidt-Hebbel (1999), who shows that 0.1-0.4 per cent of the 1.5 per cent increase in total factor productivity (TFP) in Chile in the 1980s and 1990s was attributed to pension reform. As for LCPW, one out of five countries, South Africa, shows an incorrect positive sign. For the other four countries, however, the sign is negative, consistent with our findings earlier. In other words, in these countries, growth in the capital/labour ratio accompanies a rise in economy wide productivity.

Tables 7a and 7b show the comparable results for the cointegrating vector with trend. Virtually all of the Trace and Maximum Eigenvalue tests show cointegration. The results are broadly comparable; in Table 7a we have 8 out of 11 results showing a positive effect of LPFAGDP on output per capita, while for LCPW it is 9 out of 11. In Table 7b we have 3 out of 5 with a positive effect of pension fund assets on output per head, and in 4 out of 5 cases for LCPW. Note that as shown in Section 7.3 below for the impulse responses, even where the cointegrating vector has a “wrong” sign, the dynamics may be such as to generate a long-term positive effect of pension assets on output.

As regards the trend, among eleven OECD countries, six show a negative coefficient, which implies (given normalization) that technological advances over time enhance economic growth. The same finding is obtained by McCoskey and Kao (1999), where six out of eight OECD countries are identified to have a positive and significant trend. Similar results are found for EMEs (Table 7b) where three out of five countries show significant trends enhancing economic growth.

To complement our country-by-country analysis, we derived the panel co-integration coefficients in Table 8 by averaging the individual coefficients from the above individual regressions. The formula for β_{panel} is as follows:

$$\beta_{panel} = \frac{\sum_{i=1}^{n} \beta_i}{n} \tag{15}$$

β_{panel} is the panel coefficient, β_i the coefficient for individual countries, and n the number of countries concerned.

T-values for the panel co-integration were calculated by following the formula,

$$t_{\beta,panel} = \frac{\sum_{i=1}^{n} t_{\beta_i}}{\sqrt{n}} \tag{16}$$

$t_{\beta,panel}$ is the panel t-values, and t_{β_i} the t-value for individual countries.

Again, our sample countries are grouped into All, OECD countries and EMEs. When estimating OECD and EMEs, we consider two scenarios, i.e. Panel 1 and Panel 2. In Panel 1, we include all countries, regardless of the signs of coefficients. T-ratios are given in brackets under the estimates of corresponding coefficients. In Panel 2, we utilise all coefficients of LPFAGDP that are negative, consistent with theory, excluding other countries.

The key finding is that in all cases, the effect of pension funds in the production function is positive and significant. For the All estimation, as revealed in the first row of Table 8, the coefficient of LPFAGDP without trend is –0.056, while that of LCPW is –0.646; both are of reasonable magnitude, consistent with our expectation. When taking the equation with trend the pension asset variable is smaller at -0.031 but still highly significant. The LCPW term is much smaller with the time trend, reflecting some of the outlying individual country results.

In Panel 1, where all estimates are included, for OECD countries, the LPFAGDP variable is smaller for OECD countries than for EMEs, with the former being -0.046 without trend and -0.004 with trend, while for EMEs the corresponding results are -0.076 and -0.090. These are consistent with the results from DOLS and Mean-Group estimation. LCPW is 0.2-0.8 except for OECD countries with trend where the variable is very small. The time trend is on average positive for EMEs but negative for OECD countries.

In Panel 2 where we only consider those countries with expected-sign pension asset elasticities in individual regressions, the panel coefficients are naturally larger at around -0.11 for OECD countries and -0.2 for EMEs. The capital stock variables are of more reasonable size, being 0.6-0.8 in each case.

In all, our co-integration estimations in this section, split into without and with trends, support the positive and long run relationship between growth, pension assets, capital stock and technological advances. In addition, there is further evidence that the beneficial impact of pension growth on growth is higher in EMEs than in OECD countries.

7.3 Impulse responses and variance decomposition

We now move on to impulse response tests derived from the Vector-Error-Correction Model underlying the Johansen results. The underlying rationale behind impulse responses is that a shock to one variable not only directly affects the variable itself, but also is transmitted to all of other endogenous variables through the dynamic structure of the VECM. In our example, it implies that pension fund assets can directly impact on output per worker, but it might also affect capital per worker, which in turn induces improvement on output. Results are based on the Pesaran and Shin (1998) generalised response approach. Technically, it constructs an orthogonal set of innovations that does not depend on the VAR ordering. The generalized impulse responses from an innovation to the j-th variable are derived by applying a variable specific Cholesky factor computed with the j-th variable at the top of the Cholesky ordering. It avoids the arbitrariness of the Cholesky ordering.

We specify a 25-year window given that it is expected that pension fund assets have a relatively long-period effect on both LOPW and LCPW, and hence a shorter period, e.g. 10 years might not be long enough to capture the long run effect of LPFAGDP. The results are summarized in Table 9. It can be seen that the results are virtually all positive over 5, 10 and 25-year horizons. Among OECD countries, the exceptions are Belgium, the Netherlands and Switzerland, small open economies, growth in which is more dependent on external factors. Also, pension funds in those countries tend to be defined benefit, which may reduce the beneficial effect on labour markets of pension funding. Only in Switzerland is the negative effect sizeable, however. With the trend we also find a negative effect for Brazil, which again is characterised by defined benefit funds. Note that in Malaysia, Germany and Canada, the positive impact arises despite a negative implied effect of pension assets in the cointegrating vector, and the opposite for the Netherlands and Belgium. This illustrates that the dynamics of the VECM can be such as to offset the sign of the long run effect for protracted periods.

Looking at the summary results at the bottom of the table, we see that, consistent with the various parameter estimates in sections 5, 6 and 7, there is on average a positive effect of pension fund growth on output per capita, and this is significantly larger in EMEs than in OECD countries, even when the three OECD countries with a negative effect are excluded. As regards time patterns, OECD country impulse responses tend to be smoother than those of EMEs. This may link to economic vulnerability, and greater sensitive to external factors, such as currency crises and policy shifts.

Turning to variance decompositions (Table 10), we see that the variance of pension fund assets explains a considerable proportion of the variance of output per capita in many of the countries concerned. There is a considerably higher proportion explained in EMEs than in OECD countries, suggestive of a greater potential contribution pension reform can have on the wider economy in these countries. Furthermore, whereas inclusion of a time trend in the VECM reduces the contribution of pension assets both in OECD countries and EMEs, the reduction is proportionately much larger for the former (from around 10% to 5%) than in the latter (from 30% to 25%).

9. Conclusion

Pension funds have been expanding and will continue such a trend in coming decades given the rapidly aging population and the transition from unfunded systems to funded systems such as the World Bank multi-pillar model. Research on the direct link between pension funds growth and economic growth, however, is quite scarce. In this paper, we first briefly reviewed the issue of whether and why pension assets and economic performance are correlated. In Section 2, a modified Cobb-Douglas production function was developed, where we included pension assets viewed as a shift factor. The underlying philosophy is that pension assets can affect economic growth indirectly via financial market development (Davis and Hu 2005; Walker and Lefort 2002), or by its economy-wide impact through corporate engagement (Clark and Hebb 2003; Davis 2002 and 2003) and giving rise to less labour market distortion following pension reforms (Disney 2003). In Section 4, results from our panel unit root tests indicated that all of our data are non-stationary but become stationary after first differencing, i.e. they are all I(1) variables.

We employed a variety of econometric techniques, all with certain advantages as well as disadvantages, to explore in the light of theory the existence and significance of the relationship between log of output per worker (dependent variable) and log capital per worker and log pension assets/GDP (independent variables). We included a trend in estimation, as well as showing results without trend, to capture other exogenous factors (such as structural reform and technical progress) that might affect the relationship between the capital/labour ratio and output per capita. As shown in the summary Table 11, pension assets/GDP were found to positively and significantly affect output in a variety of econometric specifications, consistently for both the OECD countries and EMEs, while as shown above, effects are consistently larger for EMEs than for OECD countries.

In more detail, in Section 5 we used the dynamic OLS (DOLS) model to examine the relationship between these I(1) variables. Results are encouraging, as we found a beneficial impact from pension assets growth to the output in the long run, which was significant in virtually all cases as indicated in Tables 4a and 4b. The results were robust when we used two different specifications, i.e. 1 lead/lag and 2 leads/lags and with and without trend. In Section 6, in view of cross sections' heterogeneity, we used dynamic heterogeneous models (Pesaran and Smith 1995) with an ARDL specification to investigate the average long run relations. The mean group estimator suggested a long run positive correlation between pension fund assets and output, but the values of the coefficients estimated vary between two methods. In both DOLS (with trend) and consistently in Mean-Group estimation we found evidence that EMEs benefit more from pension fund growth than OECD countries. For example, in Table 4b (DOLS with 2 leads/lags and trend), the positive effect is 0.049 for EMEs, while it is 0.012 for OECD countries.

In Section 7, by using the methodology developed by Johansen (1991 and 1995) we investigated whether our I(1) variables are co-integrated. As suggested by our theoretical model in Section 2, both pension assets and capital per worker in most cases are co-integrated with output per worker. In the last part of Section 7, we used impulse response and variance decomposition tests to provide some quantitative estimates as to how and to what extent a shock to pension assets can affect output per worker and capital per worker. Results from impulse responses tests indicated that for most countries, pension assets growth boosts both capital and output during the initial few years before following a gradual decline. With the exception of some small open economies with defined benefit funds, the effect of a rise in pension fund assets on economic growth is positive, validating our theoretical analysis in Section 2. Both impulse responses and variance decompositions show a larger impact or level of explanation of pension assets for EMEs. Consistent with this, the panel estimates calculated from individual VECM regressions show that the beneficial effect of pension assets growth on economic development is stronger for EMEs than OECD countries.

The overall policy implication of this research favours pension funding as a response to the challenge of ageing, as it indeed appears to offer an additional benefit to the economy in terms of productive efficiency. That said, it should be cautioned that not all countries have the necessary administrative and organizational infrastructure to develop successful pension funds (Holzmann and Hinz 2005) so careful preparation is necessary before launching such a pension reform.

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Table 1 Total assets of pension funds within 18 advanced OECD countries (as of 2000)

	Country Name	**Total assets (US$ mn)**	**As % of GDP**	**As % of Total**
AUS	Australia	188892.83	48.63	1.54
AUT	Austria	7300.00	3.87	0.06
BEL	Belgium	14400.00	5.74	0.12
CAN	Canada	310500.00	43.94	2.54
CHE*	Switzerland	268600.00	124.25	2.19
DEU	Germany	62200.00	3.33	0.51
DNK	Denmark	40100	23.05	0.33
ESP**	Spain	32806.00	5.85	0.27
GBR**	UK	1141830.72	79.87	9.33
ISL	Iceland	6700.00	78.91	0.05
ITA	Italy	48100.00	4.48	0.39
JPN	Japan	2893319.29	60.72	23.63
NLD	Netherlands	550935.92	149.09	4.50
NOR**	Norway	11300.00	7.36	0.09
NZL***	New Zealand	615.00	0.69	0.01
PRT	Portugal	12400.00	11.70	0.10
SWE	Sweden	93922.37	41.01	0.77
USA	US	6559771.48	66.87	53.58
	Total assets within OECD countries	12243693.61	42.19****	100.00

Source: See Section 3 for details. * 1998 data, ** 1999 data and ***2002 data. **** average of pension assets of GDP within OECD countries.

Table 2: Total assets of pension funds within 29 EMEs (as of 2002)

	Country Name	**Total assets (US$ mn)**	**As % of GDP**	**As % of Total**
ARG	Argentina	11409	11.16	4.05

BGR	Bulgaria	41.94	0.27	0.01
BOL	Bolivia	1144	14.9	0.41
BRA	Brazil	47656	10.53	16.92
CHL	Chile	35500	55.34	12.60
COL	Colombia	5482	6.67	1.95
CRI	Costa Rica	136	0.81	0.05
DOM	Dominican Republic	184.49	0.87	0.07
ECU	Ecuador	14.27	0.06	0.01
FJI	Fiji	846.95	45.11	0.30
HND	Honduras	3.28	0.05	0.00
HUN	Hungary	1835	2.79	0.65
IDN	Indonesia	278.21	0.05	0.10
KAZ	Kazakhstan	1432	5.92	0.51
KOR*	Korea	11500	2.49	4.08
LKA*	Sri Lanka	2697.99	16.55	0.96
MEX	Mexico	31748	4.98	11.27
MYS	Malaysia	53605.11	56.33	19.03
PAK	Pakistan	947.98	1.57	0.34
PAN	Panama	464	3.77	0.16
PER	Peru	4527	7.96	1.61
PHL	Philippines	3062.5	3.97	1.09
POL	Poland	6674	3.56	2.37
RUS	Russia	1612.7	0.47	0.57
SGP	Singapore	55526.98	63.85	19.71
SLV	Slovakia	1088	7.62	0.39
UKR	Ukraine	2.62	0.01	0.00
URG	Uruguay	893	7.25	0.32
ZAF**	South Africa	1423.63	0.01	0.51
	Total assets within EMEs	281736.65	11.55***	100.00

Source: various sources, including OECD Institutional Investors (2003), Davis and Steil (2001) and national sources. See Section 3 for details. All data are converted into and measured at US Dollars, for the convenience of across-country comparison.

* 2000 data. **, 1997 data, *** average of pension assets of GDP within EMEs.

Table 3. Panel unit root test (38 countries, 20EMEs+18OECD)

Variable	Level			1st difference			2nd difference
	IPS (2003)	LLC (2002)	Hadri (2000)	IPS (2003)	LLC (2002)	Hadri (2000)	Hadri (2000)
PFAGDP	9.37	9.21	20.76***	-14.86***	-18.42***	2.70***	1.24
p-value	1.00	1.00	0.00	0.00	0.00	0.00	0.11
OPW	8.09	3.11	26.30***	-21.99***	-22.07***	5.22***	0.88
p-value	1.00	1.00	0.00	0.00	0.00	0.00	0.19
CPW	12.10	4.49	24.28***	-4.11***	-2.08**	8.17***	-0.46
p-value	1.00	1.00	0.00	0.00	0.02	0.00	0.68

PFAGDP: Pension fund assets/GDP. OPW: Output per worker. CPW: Capital stock per worker. Panel unit root tests are based on Im, Pesaran and Shin (2003), Levin, Lin and Chu (2002) and Hadri (2000). The null hypothesis of IPS and LLC is non-stationarity, while that of Hadri is stationarity. *** significance at 1%. ** significance at 5%.

Table 4a. Estimates from dynamic OLS (DOLS) estimations (1 lead and 1 lag). Dependent variable – LOPW (37 countries)

Estimated by panel fixed effects GLS with cross section weights.

	All		OECD		EMEs	
	No trend	With trend	No trend	With trend	No trend	With trend
time trend		0.008***		0.010***		0.006***
LPFAGDP	0.047***	0.001	0.065***	0.015***	0.036***	0.019***
LCPW	0.707***	0.549***	0.662***	0.385***	0.71***	0.624***
Adjusted R-squared	0.999	0.999	0.98	0.99	0.997	0.998
S.E. of regression	0.051	0.042	0.046	0.031	0.056	0.053
OBS	570	570	383	383	187	187
No. of countries	37	37	18	18	19	19
Memo: LPFAGDP with SUR estimation			0.074***	0.021***		

Key: LPFAGDP: log of Pension fund assets/GDP. LOPW: log of output per worker. LCPW: log of capital stock per worker. ***, significant at 1%, ** significant at 5%, * significant at 10%. OBS, number of observations.

Table 4b. Estimates from dynamic OLS (DOLS) estimations (2 leads and 2 lags). Dependent variable – LOPW (33 countries)

Estimated by panel fixed effects GLS with cross section weights.

	All		OECD		EMEs	
	No trend	With trend	No trend	With trend	No trend	With trend
time trend		0.008***		0.011***		0.005***
LPFAGDP	0.051***	-0.001	0.068***	0.012**	0.058***	0.049***
LCPW	0.704***	0.542***	0.650***	0.375***	0.713***	0.653***
Adjusted R-squared	0.997	0.997	0.982	0.991	0.998	0.998
S.E. of regression	0.048	0.04	0.043	0.029	0.053	0.05
OBS	498	498	347	347	150	150
No. of countries	33	33	18	18	15	15
Memo: LPFAGDP with SUR estimation			0.078***	0.016***		

Key: LPFAGDP: log of Pension fund assets/GDP. LOPW: log of output per worker. LCPW: log of capital stock per worker. ***, significant at 1%, ** significant at 5%, * significant at 10%. OBS, number of observations.

Table 5a. Heterogeneous panel estimates of mean long run output per worker (LOPW) elasticities. (16 countries, 11OECD + 5EMEs), with trend.

	Method 1*		**Method 2***	
	LPFAGDP(η)	LCPW (ξ)	LPFAGDP(η)	LCPW(ξ)
All	0.120***	1.025***	0.08***	1.0***
OECD	0.009	0.936***	0.012	0.947***
EMEs	0.453***	1.284***	0.311***	1.189***

Key: see Table 3. Method 1 is the average of long run elasticities across countries, while method 2 is long runs from means of short run elasticities. Both methods are based on Pesaran and Smith (1995), calculation of t values based on Smith and Fuertes (2004). See Section 6 in text for details.

Table 5b. Heterogeneous panel estimates of mean long run output per worker (LOPW) elasticities. (10 countries, 7OECD + 3EMEs), with trend

	Method 1*		**Method 2***	
	LPFAGDP(η)	LCPW (ξ)	LPFAGDP(η)	LCPW(ξ)
All	0.204***	1.083***	0.11***	1.01***
OECD	0.073***	0.953***	0.047***	0.95***
EMEs	0.531*	1.38***	0.32*	1.22***

Key: see Table 3. Method 1 is the average of long run elasticities across countries, while method 2 is long runs from means of short run elasticities. Both methods are based on Pesaran and Smith (1995) , calculation of t values based on Smith and Fuertes (2004). See Section 6 in text for details.

Table 5c. Heterogeneous panel estimates of mean long run output per worker (LOPW) elasticities. (16 countries, 11OECD + 5EMEs), without trend

	Method 1*		**Method 2***	
	LPFAGDP(η)	LCPW (ξ)	LPFAGDP(η)	LCPW(ξ)
All	0.137***	0.509***	0.094***	0.549***
OECD	0.129**	0.635***	0.061**	0.724***
EMEs	0.189***	0.087	0.167***	0.167

Key: see Table 3. Method 1 is the average of long run elasticities across countries, while method 2 is long runs from means of short run elasticities. Both methods are based on Pesaran and Smith (1995) , calculation of t values based on Smith and Fuertes (2004). See Section 6 in text for details.

Table 6a. Co-integrating coefficients vector without trend; normalised on LOPW. OECD countries

	LOPW	LPFAGDP	LCPW	Trace	Max-Eigenvalue
Australia	1	-0.23***	-0.02	36.49	25.02
		[7.67]	[0.16]		
Belgium	1	-0.005	-0.68***	30.4	14.4
		[0.17]	[6.42]		
Canada	1	0.29***	-1.25***	43.3	26.2
		[4.83]	[11.36]		
Denmark	1	-0.11***	-0.76***	44.0	27.93
		[13.16]	[17.18]		
Germany	1	-0.27***	-0.53***	33.2	17.5
		[11.33]	[4.56]		
Japan	1	-0.12***	-0.56***	41.0	26.24
		[7.22]	[16.62]		
Netherlands	1	-0.11**	-0.67***	27.42	23.51
		[3.8]	[3.3]		
Sweden	1	-0.04*	-1.21***	30.02	25.24
		[1.62]	[42.14]		
Switzerland	1	0.18***	-1.23***	35.26	22.09
		[3.60]	[9.46]		
UK	1	-0.06***	-0.78***	33.65	25.10
		[12.49]	[39.13]		
USA	1	-0.04	-0.75***	31.54	22.62
		[1.49]	[17.47]		

Key: see Table 3. Co-integration estimation is based on Johansen methodology (1991 and 1995). T values are under estimates of corresponding coefficients. Lag length is selected based on a range of criteria statistics, e.g. AIC (Akaike information criterion) and SC (Schwarz information criterion). T-values are in square brackets. Under both Trace and Max-eigenvalue statistics, all countries indicate a co-integration relationship at 5% or 10% level; the only exceptions are Belgium and Germany under Max-eigenvalue statistics.

Table 6b. Co-integrating coefficients vector without trend; normalised on LOPW. Emerging market economies (EMEs)

	LOPW	LPFAGDP	LCPW	Trace	Max-Eigenvalue
Brazil	1	-0.07***	-0.12	35.70	25.20
		[13.20]	[0.36]		
Chile	1	-0.14***	-0.48***	49.60	34.80
		[23.00]	[35.82]		
Korea	1	-0.27***	-0.71***	45.1	25.8
		[5.88]	[28.18]		
Malaysia	1	0.23	-0.8***	35.4	24.9
		[5.75]	[26.67]		
South Africa	1	-0.14***	0.19***	105.6	60.9
		[17.20]	[-3.80]		

Key: see Table 3. Under both Trace and Max-Eigenvalue statistics, all countries indicate a co-integration relationship at 5% or 10% level.

Table 7a. Co-integrating coefficients vector with trend; normalised on LOPW; OECD countries

	LOPW	LPFAGDP	LCPW	Trend	Trace	Max-Eigenvalue
Australia	1	-0.23***	-0.20	0.00	44.19	25.52
		[7.67]	[0.56]	[0.30]		
Belgium	1	-0.02***	-0.03	-0.02***	74.30	50.30
		[6.05]	[0.98]	[27.68]		
Canada	1	0.26***	-0.67***	-0.01***	27.02	20.39
		[8.67]	[6.09]	[5.50]		
Denmark	1	-0.14***	-0.95***	0.004	29.35	23.77
		[4.39]	[8.04]	[-1.32]		
Germany	1	0.50***	2.87***	-0.07***	42.03	21.74
		[-3.30]	[-2.96]	[4.57]		
Japan	1	-0.12***	-0.73***	0.01	46.34	27.77
		[5.79]	[5.40]	[-1.45]		
Netherlands	1	-0.14*	-0.72***	0.002	46.34	27.77
		[1.66]	[3.16]	[-0.29]		
Sweden	1	-0.03	4.08***	-0.07***	71.27	47.82
		[0.65]	[-7.97]	[10.44]		
Switzerland	1	0.15***	-0.66**	-0.017*	47.0	23.4
		[3.42]	[2.46]	[1.66]		
UK	1	-0.16***	-1.69***	0.025***	28.20	20.83
		[5.5]	[6.6]	[-3.8]		
USA	1	-0.11**	-1.48***	-0.02*	42.01	23.14
		[2.20]	[3.08]	[1.62]		

Key: see Table 3. Under both Trace and Max-eigenvalue statistics, all countries indicate a co-integration relationship at 5% or 10% level; the only exception is Germany under Max-eigenvalue statistics.

Table 7b. Co- integrating coefficients vector with trend; normalised on LOPW; Emerging market economies (EMEs)

	LOPW	LPFAGDP	LCPW	Trend	Trace	Max-Eigenvalue
Brazil	1	0.16***	0.07***	-0.03***	84.50	49.10
		[20.33]	[14.11]	[-24.55]		
Chile	1	-0.31***	-0.70***	0.04***	65.20	41.70
		[10.33]	[10.00]	[4.11]		
Korea	1	-0.40***	-0.95**	0.02	33.62	23.94
		[4.15]	[2.37]	[-0.61]		
Malaysia	1	0.23***	-0.69***	-0.007***	40.4	25.1
		[-6.3]	[8.2]	[6.07]		
South Africa	1	-0.13***	-0.24***	-0.01***	137.1	83.9
		[29.89]	[5.59]	[10.92]		

Key: see Table 3. * Under both Trace and Max-eigenvalue statistics, all countries indicate a co-integration relationship at 5% or 10% level.

Table 8. Panel estimation of co-integrating coefficients normalized on LOPW

		Without trend		With trend		
		LPFAGDP	**LCPW**	**LPFAGDP**	**LCPW**	**T**
All	OECD+EMEs	-0.056***	-0.646***	-0.031***	-0.082***	-0.008***
		(21.6)	(64.7)	(6.2)	(4.8)	(18.7)
Panel 1	OECD	-0.046***	-0.766***	-0.004*	-0.016***	-0.013***
		(14.9)	(49.4)	(1.9)	(7.2)	(12.9)
	EMEs	-0.076***	-0.382***	-0.090***	-0.226***	0.005***
		(16.7)	(42.4)	(8.3)	(2.1)	(14.3)
Panel 2	OECD	-0.109***	-0.661***	-0.119***	-0.887***	-0.026***
		(19.2)	(47.7)	(7.8)	(12.1)	(17.4)
	EMEs	-0.153***	-0.662***	-0.280***	-0.630***	0.018***
		(21.4)	(42.4)	(25.5)	(10.4)	(3.5)

Key: see Table 3. Panel coefficients and t-values are calculated using individual estimates (see Text for details). Panel 1 includes all countries in relevant groups, while Panel 2 includes only those countries whereby estimates of coefficients are negative in individual normalized cointegrating equations.

Table 9. Summary of impulse responses of output per worker to pension assets

Percent response to one standard deviation change

	Without trend			With trend		
Years	5	10	25	5	10	25
Australia	1.13	1.15	1.14	1.10	1.12	1.11
Belgium	-0.12	-0.13	-0.20	-0.04	0.09	0.05
Canada	0.64	0.24	0.26	0.05	-0.06	-0.09
Denmark	0.24	0.43	0.42	-0.10	0.39	0.33
Germany	0.71	1.67	1.87	0.40	0.86	1.45
Japan	1.24	0.89	0.96	0.94	0.76	0.79
Netherlands	-0.28	-0.31	-0.41	-0.32	-0.36	-0.47
Sweden	1.23	1.05	1.05	0.56	1.31	0.84
Switzerland	-1.13	-0.98	-1.00	-1.13	-0.98	-1.00
UK	0.25	0.46	0.43	0.23	0.61	0.55
USA	-0.04	0.07	0.08	0.10	0.01	0.22
Brazil	1.52	-0.21	0.48	-0.40	-0.45	-0.28
Chile	1.26	2.38	1.80	3.71	5.23	2.66
Korea	3.43	4.37	4.02	4.23	4.82	4.41
Malaysia	2.45	0.69	0.63	1.93	0.43	0.25
South Africa	4.10	3.82	3.87	4.01	4.15	4.28
All	**1.04**	**0.97**	**0.96**	**0.95**	**1.12**	**0.94**
OECD	**0.35**	**0.41**	**0.42**	**0.16**	**0.34**	**0.34**
OECD positive	**0.68**	**0.75**	**0.78**	**0.41**	**0.62**	**0.65**
EME	**2.55**	**2.21**	**2.16**	**2.70**	**2.84**	**2.26**

Note: Based on Pesaran and Shin (1998) generalised response approach. See text for details

Table 10: Variance decomposition: pension fund assets on LOPW

Percent	Without trend			With trend		
Years	5	10	25	5	10	25
Australia	4.71	4.58	4.53	4.48	4.47	4.50
Belgium	9.03	7.09	5.72	5.09	7.39	11.99
Canada	8.07	5.00	2.53	1.91	1.72	0.88
Denmark	5.47	8.14	11.95	2.38	4.13	7.79
Germany	0.91	21.74	33.37	10.44	3.73	2.30
Japan	5.74	5.04	4.35	4.09	3.75	3.58
Netherlands	0.63	0.33	0.18	0.56	0.34	0.39
Sweden	28.25	40.81	52.12	6.47	5.78	7.01
Switzerland	3.70	2.43	1.32	3.79	2.49	1.36
UK	11.60	16.35	24.54	7.74	9.75	13.60
USA	19.41	9.34	4.12	15.47	7.25	4.17
Brazil	8.79	8.80	12.22	2.80	2.59	1.91
Chile	34.89	48.82	54.52	25.43	30.49	28.03
Korea	26.48	22.66	20.99	29.49	25.56	23.90
Malaysia	67.34	45.80	44.19	74.42	58.06	56.66
South Africa	14.44	14.38	14.19	4.96	5.15	5.19
All	**15.59**	**16.33**	**18.18**	**12.47**	**10.79**	**10.83**
OECD	**8.86**	**10.98**	**13.16**	**5.67**	**4.62**	**5.23**
EME	**30.39**	**28.09**	**29.22**	**27.42**	**24.37**	**23.14**

Table 11. Summary of significant effects of log pension assets/GDP on LOPW

Method/specification	All	OECD	EMEs
DOLS			
1 lead/lag no trend	+	+	+
1 lead/lag with trend	**Ins**	+	+
2 lead/lag no trend	+	+	+
2 lead/lag with trend	**Ins**	+	+
Heterogeneous panel			
Method 1 all countries trend	+	**Ins**	+
Method 2 all countries trend	+	**Ins**	+
Method 1 subset trend	+	+	+
Method 2 subset trend	+	+	+
Method 1 all countries no trend	+	+	+
Method 2 all countries no trend	+	+	+
Johansen			
All without trend	+		
All with trend	+		
Panel 1 without trend		+	+
Panel 2 without trend		+	+
Panel 1 with trend		+	+
Panel 2 with trend		+	+

Note: Ins=insignificant

Appendix 1: Variable, data source and observation period. (20EMEs+18OECD)

Country	PFAGDP	Data source	OPW	CPW	Data source
Argentina	1994-2003	FIAP (2003)	1960-2002	1960-2001	WDI (2003)
Australia	1970-2003	OECD (2003), Davis & Steil (2001), Reserve bank of Australia	1960-2002	1971-2001	WDI (2003)
Austria	1993-2000	OECD (2003)	1960-2002	1971-2001	WDI (2003)
Belgium	1981-1999	OECD (2003)	1960-2002	1971-2001	WDI (2003)
Bolivia	1997-2003	FIAP (2003)	1960-2002	1965-2001	WDI (2003)
Brazil	1984-2003	FIAP (2003)	1960-2002	1970-2001	WDI (2003)
Canada	1966-2000	OECD (2003), Davis & Steil (2001)	1965-2002	1965-2001	WDI (2003)
Chile	1981-2003	FIAP (2003)	1960-2002	1960-2001	WDI (2003)
Colombia	1994-2003	FIAP (2003)	1960-2002	1960-2001	WDI (2003)
Denmark	1966-1999	OECD (2003), Davis & Steil (2001)	1960-2002	1966-2001	WDI (2003)
Ecuador	1995-2003	FIAP (2003)	1960-2002	1965-2001	WDI (2003)
Fiji	1994-2003	National Provident Fund	1960-2002	N.A.	WDI (2003)
Germany	1966-2000	OECD (2003), Davis & Steil (2001)	1971-2002	1971-2001	WDI (2003)
Hungary	1998-2003	FIAP (2003)	1960-2002	1960-2000	WDI (2003)
Iceland	1980-2000	OECD (2003)	1960-2002	1960-2001	WDI (2003)
Indonesia	1991-1996	Social Security Association	1960-2002	1979-2001	WDI (2003)
Italy	1990-2000	OECD (2003), Davis & Steil (2001)	1960-2002	1965-2001	WDI (2003)
Japan	1969-2002	OECD (2003), Davis & Steil (2001), Institute of Pension Research	1960-2002	1960-2001	WDI (2003)
Korea	1980-2000	OECD (2003)	1960-2002	1960-2002	WDI (2003)
Luxembourg	1985-1996	OECD (2003)	1960-2002	1965-2000	WDI (2003)
Malaysia	1975-2003	Bank Negara Malaysia	1960-2002	1960-2002	WDI (2003)
Mexico	1997-2003	FIAP (2003)	1960-2002	1960-2001	WDI (2003)
Netherlands	1967-2001	OECD(2003), Davis & Steil(2001)	1960-2002	1971-2001	WDI (2003)
Norway	1980-1999	OECD 2003)	1960-2002	1960-2000	WDI (2003)
Panama	1998-2002	FIAP (2003)	1960-2002	1980-2002	WDI (2003)
Peru	1993-2003	FIAP (2003)	1960-2002	1960-2001	WDI (2003)
Philippine	1985-2002	Social Security System	1960-2002	1960-2002	WDI (2003)
Poland	1999-2003	FIAP (2003)	1990-2002	1990-2002	WDI (2003)
Portugal	1989-2000	OECD (2003)	1960-2002	1971-2001	WDI (2003)
Singapore	1983-2003	Central Provident Fund	1960-2002	1965-2002	WDI (2003)
South Africa	1980-1997	South African Reserve Bank, Beck, Demirguc-Kunt and Levine (1999)	1960-2002	1960-2002	WDI (2003)
Spain	1988-2003	OECD (2003), FIAP (2003)	1960-2002	1971-2001	WDI (2003)
Sri Lanka	1989-2000	Employees and Provident Fund	1960-2002	1960-2002	WDI (2003)
Sweden	1966-2000	OECD (2003), Davis & Steil (2001)	1960-2002	1965-2001	WDI (2003)
Switzerland	1970-1998	OECD (2003), Davis & Steil	1960-2002	1965-2001	WDI (2003)

		(2001)			
UK	1964-2002	OECD (2003), Davis & Steil (2001), National Financial Statistics (2003)	1960-2002	1970-2001	WDI (2003)
Uruguay	1996-2003	FIAP (2003)	1960-2002	1960-2001	WDI (2003)
USA	1966-2000	OECD (2003), Davis & Steil (2001)	1960-2002	1960-2000	WDI (2003)

PFAGDP: Pension fund assets/GDP. OPW: Output per worker. CPW: Capital stock per worker.

References in Table:

FIAP (2003): Federación Internacional de Administradoras de Fondos de Pensiones (International Federation of Pension Fund Administrations) in Chile.

OECD (2003): OECD Institutional Investors Database.

Davis and Steil (2001): Institutional Investors, the MIT Press; Cambridge, Mass.

Beck, Demirguc-Kunt and Levine (1999): Financial Structure and Economic Development Database. World Bank.

WDI (2003): World Development Indicators, World Bank.

Appendix 2. Individual country (16) coefficients and average short-run coefficients. Dependent variable - LOPW

Variable	Coefficient	t-Statistic	Memo: long run effect	Variable	Coefficient	t-Statistic	Memo: long run effect
Australia				**Malaysia**			
TREND	-0.002	-1.244		TREND	0.003	0.515	
LOPW(-1)	0.374*	1.985		LOPW(-1)	0.854***	4.523	
LCPW	0.593***	3.360	0.946	LCPW	0.116	0.638	0.8
LPFAGDP	0.027	1.540	0.043	LPFAGDP	-0.103	-1.612	-0.7
Belgium				**Netherlands**			
TREND	0.002*	1.907		TREND	0.008***	3.543	
LOPW(-1)	0.820***	3.185		LOPW(-1)	0.324	1.609	
LCPW	0.155	0.658	0.86	LCPW	0.604***	3.333	0.89
LPFAGDP	-0.029	-1.476	-0.16	LPFAGDP	-0.045**	-2.208	-0.07
Brazil				**South Africa**			
TREND	-0.007	-1.292		TREND	0.002	0.492	
LOPW(-1)	0.640***	3.100		LOPW(-1)	0.865***	2.938	
LCPW	0.380**	2.098	1.06	LCPW	0.135	0.490	1.0
LPFAGDP	0.074*	1.794	0.20	LPFAGDP	0.030	0.626	0.22
Canada				**Sweden**			
TREND	0.000	0.016		TREND	0.000	0.070	
LOPW(-1)	0.912***	5.866		LOPW(-1)	0.917***	6.091	
LCPW	0.082	0.610	0.94	LCPW	0.080	0.565	0.97
LPFAGDP	-0.020	-0.292	-0.23	LPFAGDP	0.006	0.256	0.07
Chile				**Switzerland**			
TREND	-0.027***	-3.288		TREND	-0.0029	-1.425	
LOPW(-1)	0.868***	5.451		LOPW(-1)	0.776***	4.198	
LCPW	0.259	1.473	1.97	LCPW	0.217	1.258	
LPFAGDP	0.168***	3.713	1.28	LPFAGDP	0.034	0.682	
Denmark				**U.K.**			
TREND	-0.003*	-1.658		TREND	-0.001	-1.082	
LOPW(-1)	0.671***	3.409		LOPW(-1)	0.688***	4.038	
LCPW	0.333*	1.744	1.01	LCPW	0.304*	1.876	0.97
LPFAGDP	0.079*	1.940	0.24	LPFAGDP	0.030*	1.715	0.10
Germany				**USA**			
TREND	-0.001	-1.262		TREND	-0.003*	-1.770	
LOPW(-1)	0.767***	9.701		LOPW(-1)	0.373**	2.160	
LCPW	0.241***	3.140	1.04	LCPW	0.610***	3.666	0.97
LPFAGDP	0.066***	3.115	0.28	LPFAGDP	0.010	0.406	0.02
Japan							
TREND	0.000	0.050					
LOPW(-1)	1.026***	6.031					
LCPW	-0.024	-0.150	0.90				
LPFAGDP	-0.010	-0.882	0.38				
Korea				**Average**			
TREND	-0.016*	-1.746		TREND	-0.003***	-18.353	
LOPW(-1)	0.615**	2.183		LOPW(-1)	0.718***	218.667	
LCPW	0.429	1.491	1.11	LCPW	0.283***	88.528	1.02
LPFAGDP	0.043	1.193	0.11	LPFAGDP	0.022***	20.457	0.12

Key: see Table 3a. Average is the average short-run coefficient, rather than the long-run coefficient. Calculation of average “t” values is from Smith and Fuertes (2004)