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

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Abstract: Pension fund assets have increased markedly during recent decades, and there are signs that this trend will continue, particularly given demographic changes and the current pattern of pension reform towards funded systems. However, research on the extent to which growth in pension assets contributes directly to economic growth is quite scarce. This is surprising since superiority of funding to pay-as-you-go links notably to the question whether funding improves economic performance sufficiently to generate the resources required to meet the needs of an ageing population. In this paper, we design a modified Cobb-Douglas production function with pension assets as a shift factor. We then employ a dataset covering 38 countries to investigate the direct link between pension assets and economic growth, using a variety of appropriate econometric methods. We find positive results for both OECD countries and Emerging Market Economies (EMEs), with some 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-think their pension systems. Typically, they switch from unfunded systems, e.g. PAYG to funded systems. Given the funded nature of many new pension schemes, pension assets have increased across many countries.

Pension fund markets in OECD countries have witnessed a sizable increase in pension assets. For example, UK pension assets were equivalent to 115.6 billion US dollars in 1980 (21.5% of GDP), but in 2000 these two figures were increased to 1226.3 billion US dollars (85% of GDP) (OECD 2003). This trend was similar in most other OECD countries. Appendix 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 whole assets and Japan and the UK followed. In terms of pension assets relative to GDP, the Netherlands had the largest figure at 149% of GDP, while this figure for New Zealand was 0.69%, the smallest across OECD countries.

As regards data for emerging market economies (EMEs) shown in Appendix 2, Chile is the country which pioneered reform towards private funded pensions and its experience is often used 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 public funded Provident pension systems in the 1950s. Other countries with significant pension assets include Brazil and Mexico. Total pension assets across our selected EME countries 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 the 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 the 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 effects of funding which lead to higher economic growth, for example via positive externalities generating more efficient capital and labour markets. A third is whether 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 (2004), Davis and Hu (2004)), 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 assets 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. In Section 5, our first econometric work is conducted with the help of dynamic OLS (DOLS) model. In Section 6, we follow the dynamic heterogeneous estimation procedures designed by Pesaran and Smith (1995) to look at the average long run relations. In Section 7, we move to co-integration tests, investigating whether there is a long run relationship between pension funds and economic growth. In Section 8, we employ dynamic Generalised Method of Moments (GMM) (Arellano and Bond 1991, and Arellano and Bover 1995) to complement our results. Section 9 concludes the paper.

Summarising the results, a strong and positive relation is found with the dynamic OLS panel model between pension assets, output and capital is found. Panel cointegration coefficients using mean-group dynamic heterogeneous models again find a positive average long run relationship between pension assets and output, 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 show that a rise in pension assets boosts output per worker initially and then follow a gradual decline, but during a 25 year period, the effect remains positive. Generally, larger effects are found for EMEs than OECD countries. Last, dynamic Generalised Method of Moments (GMM) estimation complements our earlier analysis, with a similar outcome.

1. Literature review

There is evidence of a positive effect of funding on household saving, see the survey in Kohl and O'Brien (1998). As regards an impact on economic growth via national saving, as would be predicted by a standard Solow (2000) growth model, Estelle James (1996), the principal author of "Averting the Old Age Crisis" argues that one main advantage of World Bank multi-pillar model 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 (see Holzmann 1997), as well as the costs of tax subsidies to personal saving.

A key aspect of this issue is 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. Schmidt-Hebbel (1999) estimated that pension reform in Chile raised the saving rate. Given the difficulty of pinning down how the pension reform was financed in Chile, he considered three cases, i.e. fiscal contraction 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.

On the other hand, Samwick (1999), working with a panel of 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. On the other hand, cross-section evidence, based on data of 1990 and averages of 1991-1994, suggested that countries with PAYG systems had lower saving rates than other countries.

The link between pension funds and capital markets has been widely analysed in the recent literature, as reviewed in Davis and Hu (2004). For example, Davis (2004) outlines how pension reforms which introduce elements of funding can have a positive impact on financial market development, because following such pension reforms, the functions of financial markets are improved. For example, financial systems' function of managing uncertainty and controlling risk could be strengthened with pension fund growth as pension fund managers as portfolio professionals have better expertise and knowledge than individual investors.

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 issuing 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 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, Granger cause equity market capitalisation as well as value traded, while Impavido et al (2003) find a positive relationship between contractual saving assets and bond market capitalisation/GDP. On balance, the current literature on pensions suggests a positive relation between pension funds growth and financial development. 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. But are there other channels whereby pension funds may link to economic growth?

It has been argued that pension funds might improve corporate governance (Clark and Hebb 2003; Myners 2001)[1]. 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 documented, although empirical work is largely focused on the US[2]. 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, which is ignored or dismissed by most current pensions research. An exception is Davis (2002) who argues that complementary studies at the macro level are needed in that effects of governance initiatives from institutions may go wider than the "target firms" to the whole economy. For example, besides the target firm,

[1] But, the effectiveness of pension funds' positive impact on corporate governance has been challenged by Orszag (2002) and empirical works in the US (Mitchell and Hsin 1997).
[2] 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.

pension fund activism might also affect non-target firms, as 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, in one economy tends to improve performance, the overall effect might be higher economic growth and productivity for the whole economy. This is exactly the underlying rationale behind the specification of our model as shown in Section 2.

Besides the issue of corporate governance, labour market performance is relevant here. It is well known that due to the weak link between pension contributions and benefits under defined-benefit pay-as-you-go (PAYG) systems, there is a tendency towards earlier retirement and job immobility. It has been pointed out that during 1950-1970, there was a very sharp fall in participation rate 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 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 if so, the product demand falls and producers might consider reducing the demand for labour. In view of such problems, Estelle James (1998), the principal author of Averting the Old Age Crisis (World Bank 1994), has written: "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.

Looking at growth and 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 total factor productivity (TFP), it is found that improving financial market conditions following the pension funds reform significantly positively affected TFP. But this model suffers from low "t" values which might result from high 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 total factor productivity (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 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. His results in Davis (2002) for the G-7 plus Australia reveal that institutional investors tend to boost dividends payment across both Anglo-Saxon countries and Continental Europe and Japan, although the effects differ, while fixed investment is reduced in Anglo-Saxon countries, and this result is mixed for Continental Europe and Japan. He also finds no direct effect of the size of pension funds on GDP growth. Again, Davis (2004) using a dataset of 16 OECD countries and a standard Levine-specification 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), Hu (2004) 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 (20 countries). The reverse causality is weaker and notably for emerging markets there is no strong evidence that GDP growth homogenously causes pension assets. See Appendices 3a and 3b for a results summary, and Hu (2004) for details.

Taking into account the above literature review, this paper seeks to contribute to the current 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, which includes dynamic OLS estimator, the dynamic heterogeneous models designed by Pesaran and Smith (1995) to look at the average long run relations between variables, and the dynamic generalised method of moments (GMM) estimator, one of most popular econometric methodologies. Third we directly link pension assets to economic growth in a co-integration relationship and investigate the extent to which they are correlated in the long run as well as the impulse responses 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[3];
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. Arestis et al (2004) and McCoskey and Kao (1999), among others, use the similar specification, i.e. generalised Cobb-Douglas production function with relevant additional variables such as urbanization rates set as shift factors into the standard function. Technology may then be specified as follows:

[3] 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.

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

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 a) assigns each of our country sample with the country-specific intercept and time trend (allowing for heterogeneity across countries); b) 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

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

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

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

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

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

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

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

$Q^*{}_{i,t}$ is the output per worker and $K^*{}_{i,t}$ is the 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 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 the time trend (t), 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. But first we outline issues in data construction and unit root tests.

3. Data and variables

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 is very helpful, where we obtained 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 e.g. local central banks' Financial Bulletin, 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, we have data from 1960 to 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 Table 1 for details of the variables across 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 2 presents the results of the panel unit root tests. For log 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) 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 the OLS estimators. The advantages of DOLS over the FMOLS are no requirement for initial estimation and non-parametric correction. The DOLS model, used in our paper and following Stock and Watson (1993) is as follows:

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

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 the 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.

5.2 Empirical results

Results are given in Tables 3a and 3b. As noted above, it is arbitrary to choose the length of leads and lags in DOLS model, but the practice is to use 1 or 2 leads/lags (Mark and Sul 2002, and Kao et al 999). 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.

As regards the coefficient of LPFAGDP, in Table 3a where we used 1 lead/lag of the dynamic terms, all six estimates are positive, among which four are significant as expected, covering all three country groups. In Table 3b where we used 2 leads/lags, results for OECD are quite strong, positive and significant. Results for EMEs are positive, although not significant, which may link to heterogeneous behaviour. For the All estimation, one estimate is positive and statistically significant, while the other one is insignificant. This parameter measures the total or accumulative 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 3b.

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. This estimate range is an improvement, in that they are less than those from later panel estimations where the coefficients for LCPW are at around 0.9, e.g. in Tables 4a and 4b below.

Meanwhile, the time trend term tends to be positive and significant under all cases. It implies that technological advances over time improve the output. Its inclusion means the pension variable is not proxying an omitted trend. It is notable that the LPFAGDP coefficients are smaller with the trend, however. Last, the adjusted R-square ratios are quite high in all cases.

6. Dynamic heterogeneous models

In view of the issue 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 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; a) mean group estimator; b) aggregate time-series estimator; c) pooled mean group estimator; and d) cross-section estimator. But due to other approaches' limitations[4] 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 used in this section is specified as follows:

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

Equation 12 is the standard formulation of a dynamic heterogeneous panel model, consistent with Pesaran and Smith's specification, but with the consideration of saving degree of freedom we include only one lag of the dependent variables into the right hand side of the function rather than 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 regression 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. $\hat{\lambda}_i$,

[4] For example, the pooled estimator assumes that the coefficients are homogeneous across sections, an assumption which we consider too restrictive and inappropriate here.

$\hat{\varphi}_i$ 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 $\eta = N^{-1}\sum_{i=1}^{N}\eta_i$ 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

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

where

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

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 are given in Appendix 4, where we ran the regression as Equation 12 on 16 countries[5] individually, i.e. 11 advanced OECD countries and 5 EMEs. The coefficients of lag one of output LOPW(-1) measure the dynamic effects on output, while those of LCPW and LPFAGDP measure the short-run effects on output. Not surprisingly, results vary across countries. The general pattern, however, is clear. The capital per worker ratio is frequently positive, in 14 out of 16 estimates, indicating the positive impact of the capital accumulation on output. Regarding the pension assets/GDP, 12 out of 16 estimates show a positive sign, although some are insignificant. The average short-run coefficients of all explanatory variables are given in the bottom-right corner of Appendix 4. A one per cent increase in pension assets leads to an immediate rise in output by 0.014 per cent, while capital's contribution is larger at 0.295 per cent. The average lagged dependent variable is 0.7.

Further justified in our approach by the large differentials across countries as revealed in Appendix 4, we followed the approach by Pesaran and Smith (1995) to assess the long run relation between output, pension assets and capital. Results, according to the mean-group estimators are summarized in Tables 4a and 4b for the two methods set out above. As in earlier studies, we ran three separate regressions by country groupings, i.e. all 16 countries, 11 OECD countries and 5 EMEs. Table 4a presents results, based on all 16 countries, while Table 4b, based on 10 countries, excluding Chile, Japan, Malaysia, South Africa, Sweden and Switzerland. We dropped those countries, in that most of estimates (at least 3 out of 4

[5] 22 other countries were excluded due to the small number of observations.

estimates) for those countries are not significant (See Appendix 4 for details). Therefore, their presence might distort our results from the mean-group estimators. Their exclusion, however should not leads to a conclusion that pension assets have no any impact on the output in those countries, notably in Chile. One of the 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.

Results in Table 4a 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 can raise output by 0.937 per cent under method 1 and 0.948 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).

However, it is worth noting that the estimates for EMEs under both methods in Table 4a are not significant, which might be due to the presence of some outliers. In order to address this problem, we excluded those countries, and the subsequent results are presented in Table 4b. We still have the expected signs. Meanwhile, the estimates for EMEs become significant at 5 per cent level. In addition, we found that the effects of pension fund assets and the capital stock on growth are higher in EMEs than in advanced OECD countries. Under method 1 in Table 4b, the average long-run beneficial contribution to economic growth of a one percent rise in pension assets/GDP is 0.071 per cent in EMEs, compared with 0.069 per cent in OECD countries. This differential effect becomes more discernible under method 2, as the figure is 0.068 for EMEs, while it is only 0.046 for OECD countries. The coefficients of LCPW, i.e. ξ are all less than 1 in line with theory except in two cases under the EMEs estimations.

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.

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 using the methodology developed by Johansen (1991 and 1995). The specification is as follows:

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

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

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

α is the speed of adjustment from short run deviation to long run equilibrium. $\beta^{'}$ 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 Co integration test

This section presents the estimation results for the Johansen cointegration test. 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 5a, 5b and 6 give results of our first specification, i.e. without a trend, where results of individual regressions are shown in Tables 5a and 5b, while those for panel estimation in Table 6. In most cases, the Trace and Maximum-Eigenvalue statistics indicate a co-integration relationship between our variables, and the null hypothesis of non-cointegration is rejected at either the 5% or 10% level.

As shown in Table 5a, 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 these countries, pension fund growth has a statistically significant and positive relationship with output per worker, the extent of which varies from 0.005 for the Belgium to 0.27 for Germany. The lesser effect of pension funds in Belgium is not surprising in that economic growth in small open countries might be more likely to be affected by external factors. 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 the convergence of economic development among developed OECD countries. All estimates but in Canada 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 5b. All coefficient estimates for LPFAGDP 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 0.13 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, two out of five countries, e.g. Brazil and South Africa, show an incorrect positive sign. For the other three countries, however, the sign is negative, consistent with our findings earlier. In other words, in these three countries, capital stock induces economic development.

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

$$\beta_{panel} = \frac{\sum_{i=1}^{n} \beta_i}{n} \qquad (15)$$

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,

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

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

Results are given in the left part of Table 6. 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 utilise all coefficients of LPFA that are negative, consistent with theory, excluding other countries. In Panel 2, we include all countries, regardless of the signs of coefficients. T-ratios are given in brackets under the estimates of corresponding coefficients.

For the All estimation, as revealed in the first row of Table 6, the coefficient of LPFAGDP is – 0.07, while that of LCPW is – 0.39; both are negative and at reasonable magnitude, consistent with our expectation. In Panel 1 where we only consider those countries with expected pension asset elasticities in individual regressions, the panel coefficients are all negative and statistically significant, indicating a positive linkage between LOPW, LPFAGDP and LCPW. In addition, there is some indication that the positive effect of pension assets on growth is higher in EMEs than OECD countries; the coefficient of LPFAGDP is –0.17 for our EMEs, while it is –0.10 for OECD group. This 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.

Meanwhile, in Panel 2, all countries in corresponding groups, i.e. OECD and EMEs are included into our estimation. For OECD countries, results are robust and do not change too much, although the coefficient on LPFAGDP increases slightly to –0.03, implying a less positive effect, while that of LCPW is almost same. For EMEs, however, the LCPW has a positive sign, while the LPFAGDP is still negative. Still, a more significant impact of pension assets on growth for EMEs than OECD is identified in Panel 2.

The results discussed above are for the specification without the trend. Use of a trend is consistent with McCoskey and Kao (1999) where they use a time term to identify the potential beneficial effect of technological advances on growth over time. 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 another model with a trend. Results are given in Tables 7a (OECD countries) and 7b (EMEs).

In general, results are similar to those by the specification of without trend. Again, we use Trace and Maximum-eigenvalue statistics for co-integration tests. All statistics in Table 7a for OECD countries except the Maximum-Eigenvalue for the German regression, indicate a co-integration relationship between pension assets, growth and the capital stock. In terms of LPFAGDP, coefficients are all statistically significant, and frequently (eight out of eleven countries) negative, implying a positive relationship between pension assets and economic growth. Three exceptions are Canada, Germany and Switzerland. In addition, estimates for LCPW are significant, and only Australia, Germany and Sweden show contradicting results, i.e. positive sign. The last regressor is the trend. Among eleven countries, four show a positive sign. Therefore, for other countries, the negative coefficient of the trend term implies 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.

When turning to EMEs, a co-integration relationship between growth, pension assets, the capital stock and the trend is unanimously obtained under both the Trace and Maximum-eigenvalue statistics. However, results in Table 7b do not favour a strong impact of pension assets on growth, since the coefficients of LPFAGDP are negative for three countries – Brazil, Korea and South Africa, and positive for two other countries – Chile and Malaysia. The capital stock and trend term are significant, although not identical across the six EMEs. However, they are frequently negative, implying a beneficial impact of capital and technological innovation over time on economic growth.

As in the earlier specification without trend, we calculated the panel coefficients of all three regressors based on individual regressions in Tables 7a and 7b. Again, formulae 15 and 16 were used. All but one estimates are significant, as indicated in the right part of Table 6. For all countries (OECD+EMEs), LPFAGDP and LCPW have a positive sign, while T has a negative sign. In Panel 1, all estimates are satisfactory, as all are negative. In addition, our estimated results suggest again that pension growth may have a larger impact in EMEs than OECD countries; the coefficients of LPFAGDP are -0.21 and -0.18 respectively. In Panel 2, results for OECD countries are still encouraging, both LPFAGDP and LCPW show negative sign. In addition, the magnitude of LCPW is 0.35, a large improvement than that in Panel 1. For the estimation on EMEs, the time trend ratio is negative and statistically significant, while the other two are positive.

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 evidence that the beneficial impact of pension growth on growth is higher in EMEs than in OECD countries.

7.3 Impulse responses

We now move on to impulse responses 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. In this context, we did not introduce a time trend.

In Figures 1 through 16, we report how shocks to pension fund assets (LPFAGDP) affect both output per worker (LOPW)[6] shown in top panel of each figure, and capital per worker (LCPW)[7] shown in low panel of each figure. Results are estimated for 16 countries, including 11 OECD countries and 5 EMEs, which have a valid length of observations to run the regression. We specify 25 years given that it is expected pension fund assets have a relatively long-period effect on both LOPW and LCPW, and hence so a shorter period, e.g. 10 years might not be long enough to capture the long run effect of LPFAGDP.

Estimated results for OECD countries are given in Figures 1 to 11. For each country, the impacts of pension assets on growth and capital stock share a similar trend. Always, there is an initial boost to both LOPW and LCPW from pension funds LPFAGDP. After reaching its peak in years 3-10, it drops gradually and yet remains positive until the end of our specified period. The decline in pension funds' positive effect can be explained by the decreasing marginal elasticity, i.e. increases in pension assets lead to smaller and smaller increases in output and capital before arriving at the equilibrium. Note that the impulse responses are generally significantly different from zero for at least the first few years of the period shown.

The quite stable effect of pension assets on growth and capital stock after 3-10 years across our OECD countries as indicated by the smooth and level lines also corroborates our findings of co-integration earlier that there is a long run and positive relationship between our variables of interest. The only discernible exception is Switzerland, the case for which is a little different from other countries in that as indicated in top panel of Figure 9, the line which traces a one-time shock of LPFAGDP to LOPW dips at the very beginning, before returning to the zero horizontal line and remaining positive afterwards. The short run negative effect of pension assets might to some extent be consistent with our results in Tables 5a and 7a where we found a negative link between LOPW and LPFAGDP for Switzerland. In the long run, however, the effect of a rise in pension funds on both LOPW and LCPW is positive. Such marginal deviation of Switzerland might be due to the fact that growth in small open economies is more dependent on external factors.

Results for EMEs are given in Figures 12 through 16, with generally a long run positive impact of pension assets on growth and capital stock being found which is larger and significant for longer in most cases than for OECD countries. The graphs for EMEs are not as stable and smooth as those for OECD countries, perhaps reflecting the turbulent economic history of EMEs. For example, for Brazil in Figure 12, the positive effects are relatively small

[6] LOPW is denoted as LGDPCON/POPTTL, i.e. log of GDP in constant value divided by total population in figures.
[7] LCPW is denoted as LCAPSTK/POPTTL, i.e. log of capital stock divided by total population in figures.

on both LOPW and LCPW, and do not change too much over the whole period. In contrast, for Malaysia, the effects are comparatively volatile, albeit positive, as indicated by two spikes in Figure 15. As of South Africa, such effects turn into negative by moving below the zero-horizontal line during later years of our specified period. But they have tendency to return back to positive at the end of the period. The relatively unstable impact of pension assets on economic growth and the capital stock across EMEs, during the specified period, might be an indication of the heterogeneity across countries. This is particularly an issue for small economies, as most of them are associated with economic vulnerability, and more sensitive to external factors, such as currency crises and policy shifts.

8. GMM estimation

Dynamic-panel econometric models are becoming widely used by researchers. Equations for investment incorporating effects of uncertainty (Byrne and Davis 2004) and dynamic heterogeneous models for labour demand functions (Pesaran and Smith 1995), among others, are two applications. A specific dynamic panel estimation procedure, which has recently attracted increasing attention from economists and empiricists, is a dynamic-panel Generalised Method of Moments (GMM) model. In this section, we use the GMM estimation to complement our earlier findings.

8.1 Econometric issues

The GMM estimator for dynamic panel models was originally developed by Arellano and Bond (1991), and Arellano and Bover (1995). Suppose the regression of interest, is a simple AR(1) model, as follows:

$$y_{i,t} = \gamma y_{i,t-1} + \beta' x_{i,t} + \mu_i + \varepsilon_{i,t} \tag{17}$$

$y_{i,t}$ is the dependent variable, i.e. output per worker (OPW) in our study. $x_{i,t}$ is the independent variable, i.e. pension fund assets/GDP (PFAGDP) and capital per worker (CPW). μ_i is a country-specific and time-invariant effect, while $\varepsilon_{i,t}$ is the normal disturbance term. In addition, it is assumed that $E\left[\mu_i\right] = 0$, $E\left[\varepsilon_{i,t}\right] = 0$ and $E\left[\mu_i \varepsilon_{i,t}\right] = 0$ for i =1,…,N and t=2,…,T.

8.1.1 First-differenced GMM

The usual way to remove the country-specific effect, μ_i, is to take the first-difference of both sides of Equation (17), thus we obtain,

$$y_{i,t} - y_{i,t-1} = y_{i,t-1} - y_{i,t-2} + \beta'(x_{i,t} - x_{i,t-1}) + \varepsilon_{i,t} - \varepsilon_{i,t-1} \tag{18}$$

or

$$\Delta y_{i,t} = \Delta y_{i,t-1} + \beta' \Delta x_{i,t} + \Delta \varepsilon_{i,t} \tag{19}$$

If the error term $\varepsilon_{i,t}$ is not serially correlated and independent variables $x_{i,t}$ are weakly exogenous, Arellano and Bond (1991) propose the following sets of moment restrictions:

$$E\left[y_{i,t-s}\Delta\varepsilon_{i,t} \right] = 0 \quad for \quad t = 3,...T \quad and \quad s \geq 2 \tag{20}$$

$$E\left[x_{i,t-s}\Delta\varepsilon_{i,t} \right] = 0 \quad for \quad t = 3,...T \quad and \quad s \geq 2 \tag{21}$$

Given that the lagged values of explanatory variables are not correlated with the first-differences of error terms, the authors suggest the lagged levels of x and y could be used as potential instruments to estimate the first-differenced Equation (19). Specifically, the instruments available are as follows:

$$z_i = \begin{bmatrix} y_{i,1} & 0 & .. & 0 & 0 & .. & 0 & 0 & 0 & .. & 0 \\ .. & .. & .. & .. & .. & .. & .. & .. & .. & .. & .. \\ 0 & y_{i,1} & .. & y_{i,t-2} & 0 & .. & 0 & 0 & 0 & .. & 0 \\ .. & .. & .. & .. & .. & .. & ... & .. & .. & .. & .. \\ 0 & 0 & .. & 0 & y_{i,1} & .. & y_{i,t-2} & x_{i,1} & 0 & .. & 0 \\ .. & .. & .. & .. & .. & .. & .. & .. & .. & .. & .. \\ 0 & 0 & .. & 0 & 0 & .. & 0 & 0 & y_{i,1} & .. & x_{i,t-2} \end{bmatrix} \quad 22)$$

This is the central point of the standard or so-called first-differenced GMM estimator. However, the first-differenced GMM estimator, although promising, suffers from flaws. First, if times series are persistent over time, or simply I(1) variables (Bond 2002), such an estimator is inefficient in that the instruments available for the equations in the first-differences have tendency to be weak, i.e. lagged levels are weakly correlated to subsequent first-differences, the consequence of which is serious finite sample biases (Blundell and Bond 1998, 2000). This potential problem is likely to be present in our dataset, as Table 2 indicates that PFAGDP, OPW and CPW are non-stationary, i.e. persistently up-trending over time. Therefore, lagged levels of, e.g. CPW, are not suitable instruments in this context, because they are weakly correlated with $\Delta\varepsilon_{i,t}$.

Second, when running panel models, we are interested in the cross-country relationship as well as time-series relationship. However, by using first-differencing technique as in Equation (18), the time-series information remains while the pure cross-country information is lost. Such information loss in cross-country dimension is particularly serious, if they are considered to play an important role in the data generation process (DGP).

8.1.2 System GMM

In views of those shortcomings associated with first-differenced GMM estimator mentioned above, an extended GMM estimator is proposed (Arellano and Bover 1995), which is also usually known as system GMM estimator. The most innovative element of the system GMM estimator is the identification of an extra set of moment restrictions as follows:

$$E\left[\Delta y_{i,t-1}(\mu_i + \varepsilon_{i,t}) \right] = 0 \quad for \quad t = 3,...T \tag{23}$$

$$E\left[\Delta x_{i,t-1}(\mu_i + \varepsilon_{i,t}) \right] = 0 \quad for \quad t = 3,...T \tag{24}$$

Together with equations (20) and (21), system-GMM estimators entail an enlarged set of instruments. Specifically, they are shown in equation (25).

$$z_i^* = \begin{bmatrix} z_i & 0 & 0 & .. & 0 & 0 & .. & 0 & 0 & 0 & .. & 0 \\ 0 & \Delta y_{i,2} & 0 & .. & 0 & 0 & .. & 0 & 0 & 0 & .. & 0 \\ .. & .. & .. & .. & .. & .. & .. & .. & .. & .. & .. & .. \\ 0 & 0 & \Delta y_{i,2} & .. & \Delta y_{i,t-1} & .. & .. & 0 & 0 & 0 & .. & 0 \\ .. & .. & .. & .. & .. & .. & .. & .. & .. & .. & .. & .. \\ 0 & 0 & 0 & .. & 0 & \Delta y_{i,2} & .. & \Delta y_{i,t-1} & \Delta x_{i,3} & 0 & .. & 0 \\ .. & .. & .. & .. & .. & .. & .. & .. & .. & .. & .. & 0 \\ 0 & 0 & 0 & .. & 0 & 0 & .. & 0 & 0 & \Delta y_{i,2} & .. & \Delta x_{i,t-1} \end{bmatrix} \quad (25)$$

Where, z_i is the instrument set in Equation 21.

According to this approach, lagged differences of variables are used as instruments in equations for levels, in addition to using lagged levels in the equations for first differences. When equations based on levels are estimated under the system-GMM estimator, the cross-country information remains, an improvement from the differenced-GMM estimator. Meanwhile, variables in levels are more likely to be correlated with their instruments, than variables in first differences. Hence, even if our data are persistent, the estimation procedure is still efficient and consistent. By using Monte Carlo simulations, Blundell and Bond (1998) find that finite sample bias is significantly reduced, and estimation precision is improved by exploiting more moment conditions as in equation (25) compared to those in equation (22).

As we have shown above, when running both differenced and systems GMM estimators, a range of moment conditions are identified. But statistically, there is the risk of over-identifying restrictions. Therefore diagnostic tests should be used to evaluate the validity of those moment restrictions and further instruments. Two notable tests are Sargan test for first-differenced GMM estimator, and difference Sargan test for system-GMM estimator respectively. Both Sargan tests are distributed χ^2, and under null hypothesis, the instruments are valid.

In addition, the validity of the GMM estimator depends on the assumption that error terms are not serially correlated, i.e. $E\left[\varepsilon_{i,t}\varepsilon_{i,s}\right]=0$ where $i=1,...,N$ and $s \neq t$. Therefore, we need to test the null hypothesis of no serial correlation in the error terms, so as to be confident about the legitimacy of differenced and/or system GMM estimators.

8.1.3 Merits of dynamic GMM model

Before moving to our empirical results, we would like to outline several advantages of dynamic GMM models over other models.

First, dynamic GMM models allow for the identification of both long run and short run effects. As shown in Equation 17, the introduction of term $y_{i,t-1}$ indicates the inertia of the dependent variables, or the immediate effect, in response to changes in the explanatory variables. Large values of $y_{i,t-1}$, e.g. close to unity may be viewed as a signal of rapid convergence to long run effects. It is worth noting that other estimation procedures, e.g. error correction model, are also able to fulfill the same task. Recent applied work in a panel data context, among others, include Davis and Zhu (2004).

Second, the strong assumption of strict exogeneity of regressors is relaxed under the GMM estimator. Based on the GMM approach, the $x_{i,t}$ series can be endogenous, i.e. correlated with the error terms, $\varepsilon_{i,t}$ and earlier shocks $\varepsilon_{i,t-s}$ in Equation 17. In other words, we assume explanatory variables are weakly exogenous, allowing for $x_{i,t}$ to be affected by current and past, but not future realisations of $\varepsilon_{i,t}$. Therefore, the relaxation of strict exogeneity eases the worries regarding simultaneity in regressions. This worry is due to not least the presence of the lagged dependent variable on the right hand side.

Last, but not least, the measurement error problem is reduced by GMM estimators. Measurement errors arise partly when the economic indicators, however carefully designed, are not able to accurately capture the underlying. Or simply, data collected are not correct due to typing errors in data input. Mathematically, suppose $y_{i,t}$ is the true series we are interested in, but instead we observed $\hat{y}_{i,t} = y_{i,t} + m_{i,t}$, where $m_{i,t}$ is the so-called measurement error term. Bond et al (2001), however, prove that this issue is not as serious under the system-GMM estimator. For the equations in levels, normal instruments are still valid as in Equation (24) excluding Z, while for the equations in first-differences, what we need to do is only to drop the lagged levels of variables dated t-2 as instruments, i.e. use lagged levels dated t-3 and earlier.

Largely due to the merits of dynamic panel GMM estimations, which currently is believed to be the best available, researchers have applied this methodology into a range of economic contexts. Beck and Levine (2004) use it, revisiting the relationship between finance and growth; Loayza et al (2000) on the issue of private saving across the world, and Bond et al (2001) on the issue of Slow growth model.

Despite GMM estimator's advantages outlined above, however, care should be taken in interpreting GMM results in next section of this paper, in that GMM is most appropriate when N is large and T is small (Bond 2002). But as of our dataset, neither is the case; for example, we only have data covering 38 countries, while observations range from 5 years to 35 years (see Table 1 for details). Therefore, results from GMM are only considered to be complementary to our previous findings.

8.2 Empirical results

Table 8 gives the results of dynamic GMM estimations. As usual, we run three separate regressions, i.e. on all 37 countries[8], OECD countries (18) and EMEs (19). Also, we use both one-step and two-step estimations, but choose estimated results from one of them based on diagnostic tests, i.e. Sargan tests and AR tests. In addition, we present short run coefficients (SR) in row 1, and long run coefficients (LR) in row 2. LR coefficients are calculated using the formula: $\beta^* = \hat{\beta} \big/ 1-\hat{\gamma}$, where $\hat{\beta}$ is the estimated coefficient for explanatory variables in Equation 17, and $\hat{\gamma}$ is the estimated coefficient for the lagged dependent variable, i.e. $y_{i,t-1}$.

As regards the ALL regression, the signs of both LPFAGDP and LCPW are positive, as expected, while the latter estimate is marginally significant at 10 per cent, and the former

[8] Poland is excluded from our regressions, because observations for pension assets are too few.

estimate is insignificant. The Sargan test reveals the validity of our instruments chosen, as the corresponding p-value is 1. By using the formula above, we computed the long run estimates. For LPFAGDP, it is 0.058, similar as results in Tables 4a and 4b, while for the LCPW is 0.759.

When estimating for only OECD countries, the coefficient of LPFAGDP is negative and significant, while that of LCPW is positive and significant. The result for pension assets is surprising, but the estimate is not reliable in this case, as not only AR(1) but also AR(2) and AR(3) show the presence of serial correlation in the specification, although the Sargan test is satisfactory. The last regression is on EMEs, results of which are encouraging. The coefficient of LPFAGDP is 0.004, positive and significant at 1 per cent level. The result for LCPW is positive, although insignificant. The Sargan test again indicates the validity of the instruments we selected. In addition, AR(2) and AR(3) have p-values at 0.219 and 0.106 respectively, indicating no serial correlation in our EMEs regression, while AR(1) correctly is serially correlated.

9. Conclusion

Pension fund markets 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 e.g. 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 reviewed briefly the issue of whether and why pension assets and economic performance are correlated, by drawing on relevant literature. 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 2004; 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). As shown in the summary Table 9 pension assets/GDP were found to positively and significantly affect output in a variety of econometric specifications, consistently for the OECD countries but frequently also for EMEs.

In more detail, in Section 5 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 most cases as indicated in Tables 3a and 3b. The results were robust when we used two different specifications, i.e. 1 lead/lag and 2 leads/lags. 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. Meanwhile, we also found evidence that EMEs benefit more from pension fund growth than OECD countries. For example, the right panel of Table 4b shows that the positive effect is 0.068 for EMEs, while it is 0.046 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 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. Despite some variations, e.g. Germany, from the middle to the end of our specified period, however, as for all countries, the effect of a rise in pension fund assets on economic growth is positive, validating our theoretical analysis in Section 2. In addition, the panel estimates calculated from individual regressions show that the beneficial effect of pension assets growth to economic development and capital stock is stronger fro EMEs than OECD countries.

In the last section, dynamic Generalised Method of Moment (GMM) estimation, was used, partly as a robustness check. Results, however, were not strongly in favour of a positive relationship between pension assets, growth and capital except in the EMEs regression, which counts as the strongest result. As we mentioned earlier, large N and small T (Bond 2002) are the most appropriate implementing context for the GMM estimator, while they are not the case for this paper. Therefore, results are considered to be only complementary.

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

Country	PFAGDP	Data source	OPW	CPW	Data source
	Observation Period		Observation period		
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 (2001)	1960-2002	1965-2001	WDI (2003)
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. FIAP(2003): Federación Internacional de Administradoras de Fondos de Pensiones (International Federation of Pension Fund Administrations) in Chile.
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Table 2. 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 3a. Estimates from dynamic OLS (DOLS) estimations (1 lead and 1 lag). Dependent variable - LOPW

	All		OECD		EMEs	
	No trend	With trend	No trend	With trend	No trend	With trend
time trend		0.008***		0.010***		0.003**
LPFAGDP	0.038***	0.001	0.065***	0.015***	0.014***	0.010
LCPW	0.718***	0.562***	0.662***	0.385***	0.700***	0.664***
Adjusted R-squared	1.000	1.000	1.000	1.000	1.000	1.000
S.E. of regression	0.108	0.041	0.044	0.031	0.056	0.057
OBS	548	548	383	383	165	165
No. of countries	36	36	18	18	18	18

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 3b. Estimates from dynamic OLS (DOLS) estimations (2 leads and 2 lags). Dependent variable - LOPW

	All		OECD		EMEs	
	No trend	With trend	No trend	With trend	No trend	With trend
time trend		0.008***		0.011***		0.002*
LPFAGDP	0.042***	-0.005	0.068***	0.012**	0.013	0.012
LCPW	0.714***	0.555***	0.650***	0.375***	0.707***	0.678*
Adjusted R-squared	1.000	1.000	1.000	1.000	1.000	1.000
S.E. of regression	0.045	0.091	0.041	0.028	0.061	0.048
OBS	477	477	347	347	130	130
No. of countries	32	32	18	18	14	14

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 4a. Heterogeneous panel estimates of mean long run output per worker (LOPW) elasticities. (16 countries, 11OECD + 5EMEs).

	Method 1*		**Method 2***	
	LPFAGDP(η)	LCPW (ξ)	LPFAGDP(η)	LCPW(ξ)
All	0.028*	0.951***	0.048*	0.974***
OECD	0.034**	0.937***	0.023**	0.948***
EMEs	0.037	0.977***	0.140	1.072***

Key: see Table 3a. 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). See Section 6 in text for details.

Table 4b. Heterogeneous panel estimates of mean long run output per worker (LOPW) elasticities. (10 countries, 8OECD+2EMEs)

	Method 1*		**Method 2***	
	LPFAGDP(η)	LCPW (ξ)	LPFAGDP(η)	LCPW(ξ)
All	0.064***	0.968***	0.050***	0.962***
OECD	0.069***	0.953***	0.046***	0.951***
EMEs	0.071**	1.019***	0.068**	1.011***

Key: see Table 3a. 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). See Section 6 in text for details.

Table 5a. Co-integrating coefficients vector without trend; normalised on LOPW. OECD countries

	LOPW	LPFAGDP	LCPW	Test statistics	
				Trace	Max-Eigenvalue
Australia	1	-0.21***	-0.12	36.49	25.02
		[5.94]	[0.91]		
Belgium	1	-0.005	-0.68***	5.58	10.38
		[0.17]	[6.42]		
Canada	1	0.22***	-1.08***	23.67	20.26
		[-5.09]	[13.72]		
Denmark	1	-0.11***	-0.76***	3.96	27.93
		[13.16]	[17.18]		
Germany	1	-0.27***	-0.53***	4.50	17.47
		[11.33]	[4.56]		
Japan	1	-0.12***	-0.56***	5.10	26.24
		[7.22]	[16.62]		
Netherlands	1	-0.06**	-0.71***	27.42	23.51
		[2.08]	[6.35]		
Sweden	1	-0.04	-1.21***	30.02	25.24
		[1.62]	[42.14]		
Switzerland	1	0.36***	-1.71***	35.26	22.09
		[-4.20]	[7.76]		
UK	1	-0.06***	-0.78***	33.65	25.10
		[12.49]	[39.13]		
USA	1	-0.03	-0.77***	31.54	22.62
		[1.49]	[17.47]		

Key: see Table 3a. Co-integration estimation is based on Johansen methodology (1991 and 1995). Standard errors 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 5b. Co-integrating coefficients vector without trend; normalised on LOPW. Emerging market economies (EMEs)

	LOPW	LPFAGDP	LCPW	Test statistics	
				Trace	Max-Eigenvalue
Brazil	1	-0.05***	4.12***	17.13	14.56
		[5.35]	[-6.04]		
Chile	1	-0.13***	-0.44***	37.10	25.07
		[11.04]	[15.19]		
Korea	1	-0.27***	-0.71***	7.51	7.51
		[5.88]	[28.18]		
Malaysia	1	-0.27	-0.53***	10.57	18.60
		[1.58]	[3.94]		
South Africa	1	-0.14***	0.19***	8.30	8.30
		[17.20]	[-3.80]		

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

Table 6. Panel estimation of co-integrating coefficients

Without trend				With trend		
		LPFAGDP	LCPW	LPFAGDP	LCPW	T
All	OECD+EMEs	0.07***	-0.39***	0.03***	0.01*	-0.02***
		[21.82]	[52.43]	[10.94]	[7.07]	[16.57]
Panel 1	OECD	-0.10***	-0.68***	-0.18***	-1.47***	-0.02***
		[18.50]	[50.26]	[11.73]	[14.14]	[19.51]
	EMEs	-0.17***	-0.56***	-0.21***	-0.88	0.01***
		[18.36]	[16.76]	[21.27]	[1.21]	[5.30]
Panel 2	OECD	-0.03***	-0.81***	-0.05***	-0.35*	0.00***
		[13.93]	[51.94]	[5.03]	[8.60]	[13.62]
	EMEs	-0.17***	0.53***	0.21***	0.81	-0.08***
		[18.36]	[16.76]	[12.10]	[-0.1]	[9.44]

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

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

	LOPW	LPFAGDP	LCPW	Trend	Test statistics	
					Trace	Max-Eigenvalue
Australia	1	-0.22***	0.27	-0.01	44.19	25.52
		[5.89]	[-0.55]	[0.85]		
Belgium	1	-0.02***	-0.03	-0.02***	74.30	50.30
		[6.05]	[0.98]	[27.68]		
Canada	1	0.24***	-0.79***	-0.01***	27.02	20.39
		[-9.71]	[10.40]	[4.62]		
Denmark	1	-0.14***	-0.95***	0.00	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	-0.12	-0.73***	0.01	46.34	27.77
		[1.27]	[3.16]	[-0.29]		
Sweden	1	-0.03	4.08***	-0.07***	71.27	47.82
		[0.65]	[-7.97]	[10.44]		
Switzerland	1	0.10***	-0.12	-0.01***	67.55	20.63
		[-3.47]	[0.87]	[8.23]		
UK	1	-0.77***	-6.57***	0.16***	28.20	20.83
		[7.57]	[8.27]	[-7.24]		
USA	1	-0.06	-1.13***	0.01	42.01	23.14
		[1.56]	[2.85]	[-0.92]		

Key: see Table 3a. 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-eigenvale statistics.

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

					Test statistics	
	LOPW	**LPFAGDP**	**LCPW**	**T**	**Trace**	**Max-Eigenvalue**
Brazil	1	-0.09**	3.83***	0.00	31.29	19.99
		[2.81]	[-5.86]	[-1.13]		
Chile	1	1.54***	1.46**	-0.37***	61.65	36.09
		[-6.23]	[-2.65]	[5.85]		
Korea	1	-0.40***	-0.95**	0.02	33.62	23.94
		[4.15]	[2.37]	[-0.61]		
Malaysia	1	0.13***	-0.03	-0.04***	14.27	24.46
		[-3.56]	[0.32]	[6.07]		
South Africa	1	-0.13***	-0.24***	-0.01***	14.71	14.71
		[29.89]	[5.59]	[10.92]		

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

Table 8. Estimates from dynamic GMM estimations. Dependent variable - LOPW

	All			OECD			EMEs		
	LOPW(-1)	LPFAGDP	LCPW	LOPW(-1)	LPFAGDP	LCPW	LOPW(-1)	LPFAGDP	LCPW
Coefficient (SR)	0.984***	0.001	0.012	0.462***	-0.065***	0.492***	0.999***	0.004*	0.006
Coefficient (LR)		0.058	0.759		-0.121	0.915		3.810	5.240
P-value	[0.000]	[0.442]	[0.109]	[0.000]	[0.000]	[0.000]	[0.000]	[0.093]	[0.752]
No. of countries	37			18			19		
OBS	660			438			222		
Sargan test	2.68E+07***			3.12E-15***			3.000E+03***		
P-value	[1.000]			[1.000]			[1.000]		
AR(1)	3.82***			3.32***			2.783**		
P-value	[0.000]			[0.001]			[0.005]		
AR(2)	2.515**			3.237***			1.228		
P-value	[0.012]			[0.001]			[0.219]		
AR(3)	2.297**			3.22***			1.616		
P-value	[0.022]			[0.001]			[0.106]		

Key: see Table 3a. Lag one of LOPW and dummy variables for time trend were also included in our regressions, but not reported. Wald is Wald test for joint significance. Sargan test is to test the null hypothesis of valid instruments. AR is to test the null hypothesis of no serial correlation. SR, short-run, and LR, long-run.

Table 9. 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**	**+**	**Ins**
2 lead/lag no trend	**+**	**+**	**Ins**
2 lead/lag with trend	**Ins**	**+**	**Ins**
Heterogeneous panel			
Method 1 all countries	**+**	**+**	**Ins**
Method 2 all countries	**+**	**+**	**Ins**
Method 1 subset	**+**	**+**	**+**
Method 2 subset	**+**	**+**	**+**
Johansen			
All	**(-)**	**na**	**Na**
Panel 1 without trend	**Na**	**+**	**+**
Panel 2 without trend	**Na**	**+**	**+**
Panel 1 with trend	**Na**	**+**	**+**
Panel 2 with trend	**Na**	**+**	**(-)**
GMM			
Dynamic estimates	**Ins**	**(-)**	**+**

Note: Ins=insignificant

Appendix 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.

Appendix 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.

Appendix 3a. F-Test results of Panel Granger causality estimation (Pension assets to GDP growth)

Lag	All countries		OECD countries		EMEs	
	HONC	HONC	HONC	HOC	HONC	HOC
1	0.00	443.89***	0.00	469.20***	0.00	1546.26***
2	8.16***	525.19***	0.04	275.52***	0.05	1077.20***
3	9.28***	192.59***	0.30	185.86***	9.97***	271.90***
4	0.12	493.25***	0.45	167.99***	-3.75***	397.33***
5	0.39	479.08***	0.79	200.39***	0.25	408.12***
	No	OBS	No	OBS	No	OBS
1	38	460	18	312	20	148
2	35	379	18	262	17	117
3	32	310	18	217	14	93
4	28	254	18	177	10	77
5	25	208	18	142	7	66

HONC is homogeneous non-causality hypothesis and HOC is homogenous causality hypothesis. *** Indicates rejection at 1%. No, number of countries, and OBS, observation.

Appendix 3b. F-Test results of Panel Granger causality estimation (GDP growth to pension assets)

Lag	All countries		OECD countries		EMEs	
	HONC	HOC	HONC	HOC	HONC	HOC
1	HO	HONC	HO	HONC	HO	HONC
2	0.00	2223.27***	0.000	176.63***	0.00	N.A.
3	0.58	1359.97***	7.847***	21.90***	-9.50***	N.A.
4	1.19	897.62***	0.036	196.49***	-2.89***	N.A.
5	-0.24	776.73***	0.079	223.77***	1.26**	N.A.
	-1.55*	482.01***	1.955***	134.98***	2.27***	N.A.
1	No	OBS	No	OBS	No	No
2	38	512	18	336	20	N.A.
3	38	426	18	285	20	N.A.
4	35	346	18	235	17	N.A.
5	30	278	18	190	12	N.A.

HONC is homogeneous non-causality hypothesis and HOC is homogenous causality hypothesis. *** Indicates rejection at 1%, * rejection at 5%. No, number of countries, and OBS, observation.

Appendix 4. Individual country (16) coefficients and average short-run coefficients. Dependent variable - LOPW

Variable	Coefficient	t-Statistic		Variable	Coefficient	t-Statistic
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		LCPW	0.116	0.638
LPFAGDP	0.027	1.540		LPFAGDP	-0.103	-1.612
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		LCPW	0.604***	3.333
LPFAGDP	-0.029	-1.476		LPFAGDP	-0.045**	-2.208
Brazil				**South Africa**		
TREND	0.001**	2.279		TREND	0.002	0.492
LOPW(-1)	0.544**	2.531		LOPW(-1)	0.865***	2.938
LCPW	0.421	2.189		LCPW	0.135	0.490
LPFAGDP	0.014	1.310		LPFAGDP	0.030	0.626
Canada				**Sweden**		
TREND	-0.002	-1.244		TREND	0.000	0.070
LOPW(-1)	0.374*	1.985		LOPW(-1)	0.917***	6.091
LCPW	0.593***	3.360		LCPW	0.080	0.565
LPFAGDP	0.027	1.540		LPFAGDP	0.006	0.256
Chile				**Switzerland**		
TREND	-0.005	-0.732		TREND	-0.003	-1.305
LOPW(-1)	1.117***	6.067		LOPW(-1)	0.770***	3.879
LCPW	-0.082	-0.435		LCPW	0.224	1.207
LPFAGDP	0.036	1.108		LPFAGDP	0.032	0.601
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		LCPW	0.304*	1.876
LPFAGDP	0.079*	1.940		LPFAGDP	0.030*	1.715
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		LCPW	0.610***	3.666
LPFAGDP	0.066***	3.115		LPFAGDP	0.010	0.406
Japan						
TREND	0.000	0.050				
LOPW(-1)	1.026***	6.031				
LCPW	-0.024	-0.150				
LPFAGDP	-0.010	-0.882				
Korea				**Average**		
TREND	-0.016*	-1.746		TREND	-0.001***	-8.189
LOPW(-1)	0.615**	2.183		LOPW(-1)	0.698***	217.415
LCPW	0.429	1.491		LCPW	0.295***	95.560
LPFAGDP	0.043	1.193		LPFAGDP	0.014***	13.753

Key: see Table 3a. Average is the average short-run coefficient, rather than the long-run coefficient.

Figure 1. Impulse responses. Australia

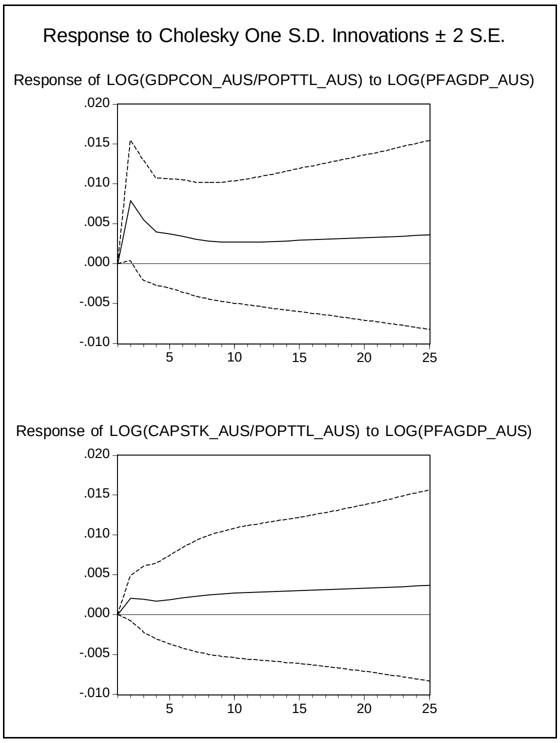

Figure 2. Impulse responses. Belgium

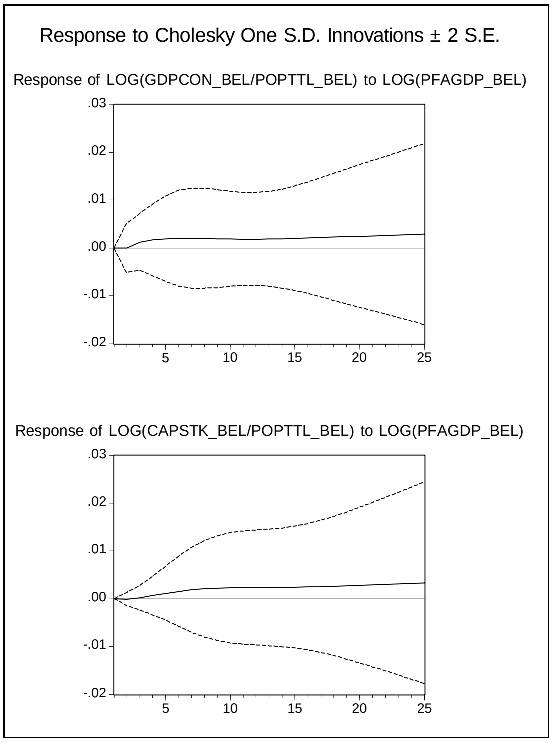

Figure 3. Impulse responses. Canada

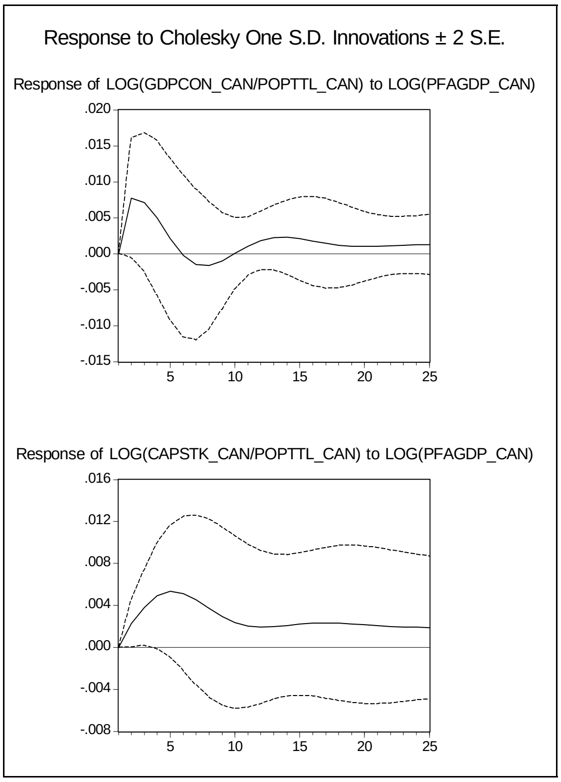

Figure 4. Impulse responses. Denmark

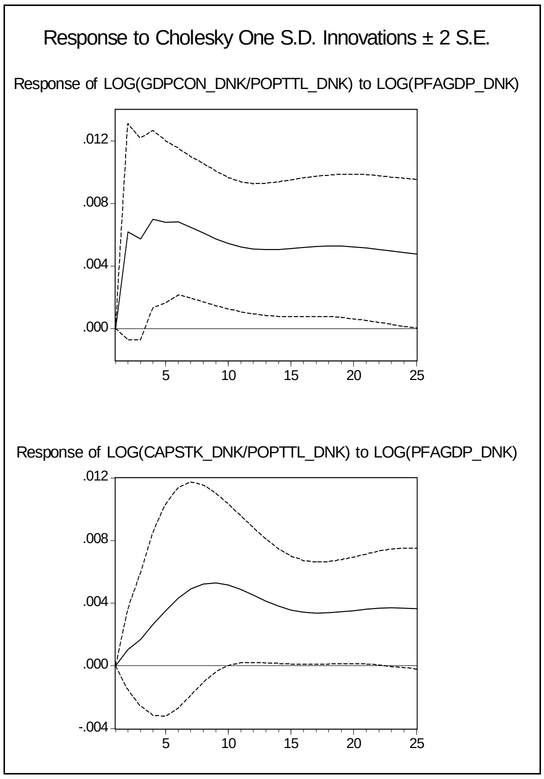

Figure 5. Impulse responses. Germany

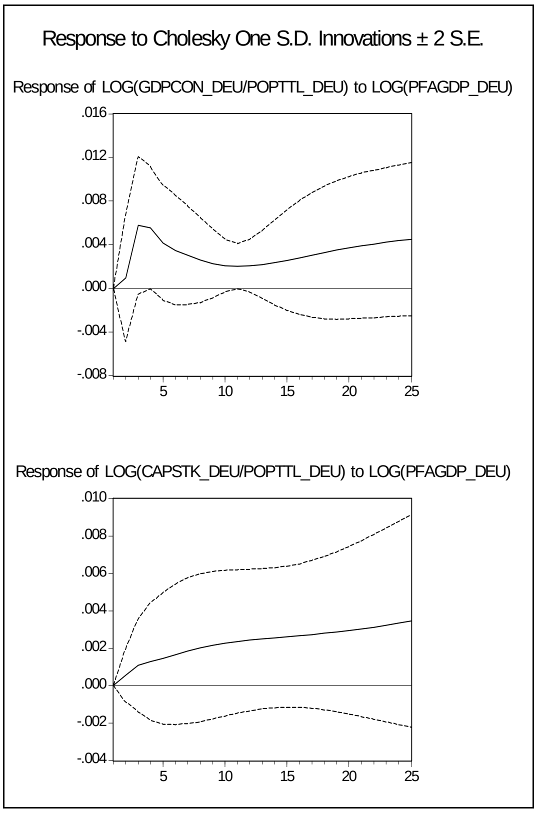

Figure 6. Impulse responses. Japan

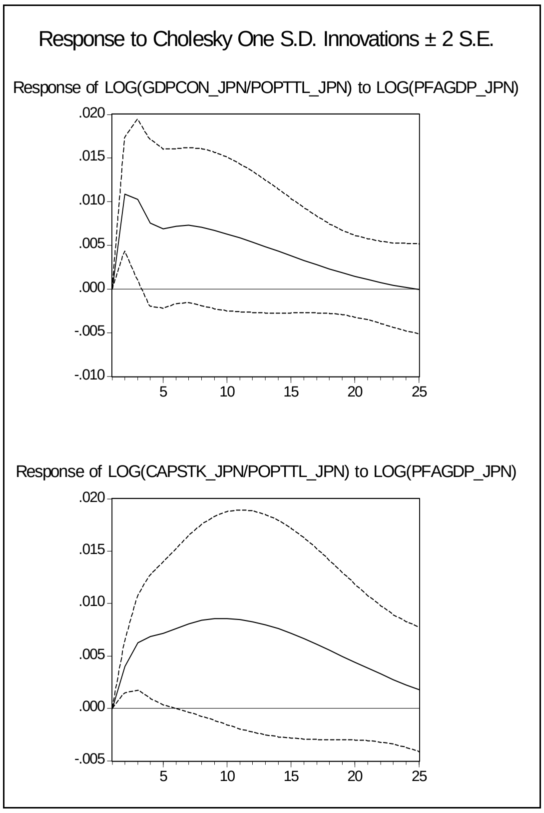

Figure 7. Impulse responses. Netherlands

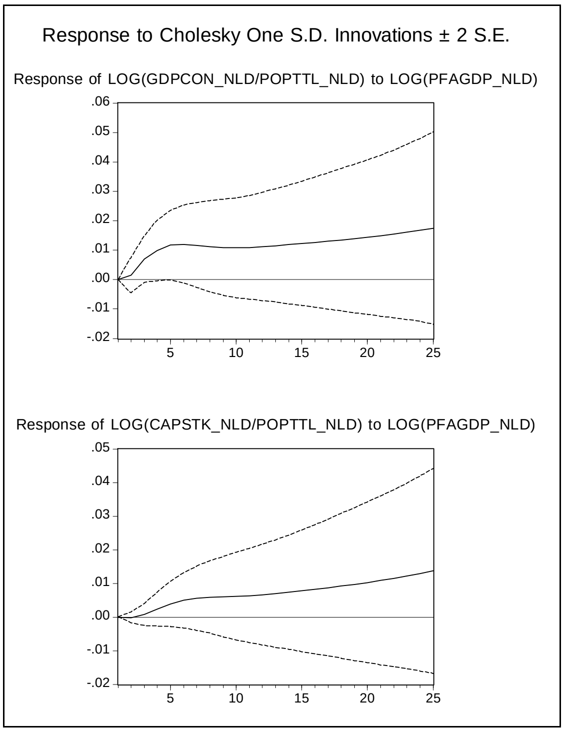

Figure 8. Impulse responses. Sweden

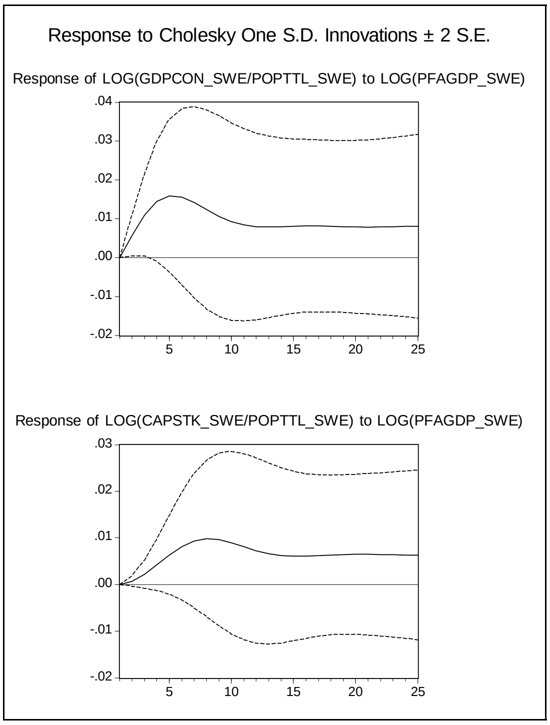

Figure 9. Impulse responses. Switzerland

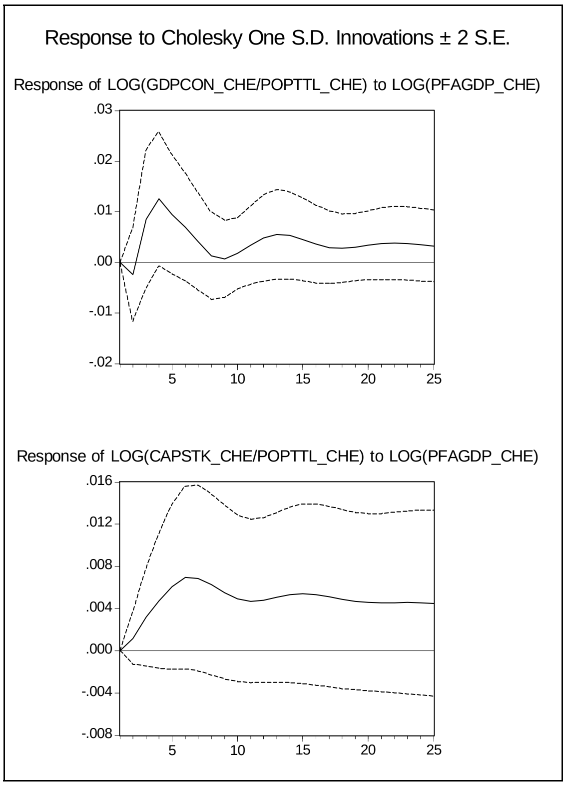

Figure 10. Impulse responses. U.K.

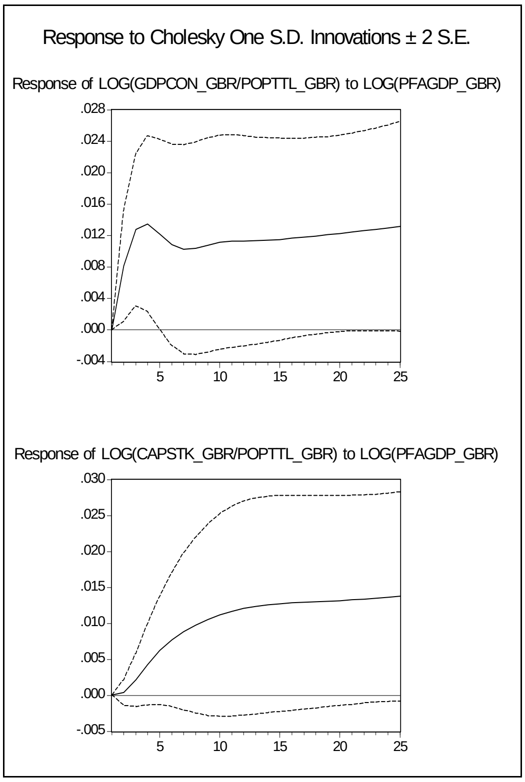

Figure 11. Impulse responses. U.S.A.

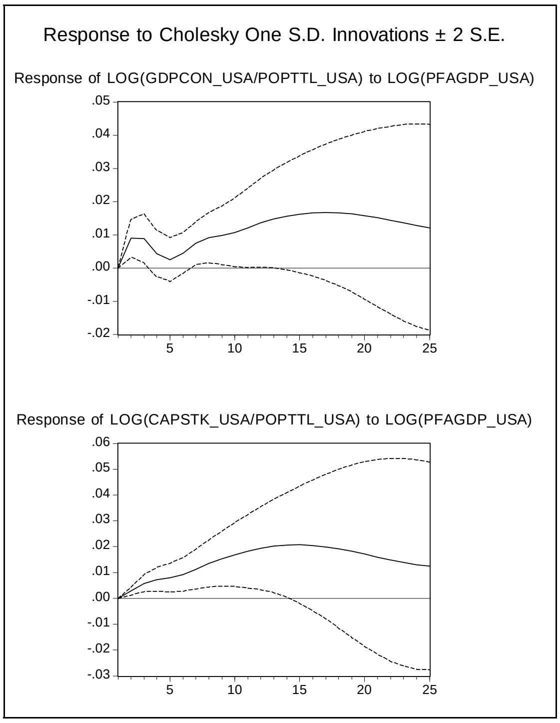

Figure 12. Impulse responses. Brazil

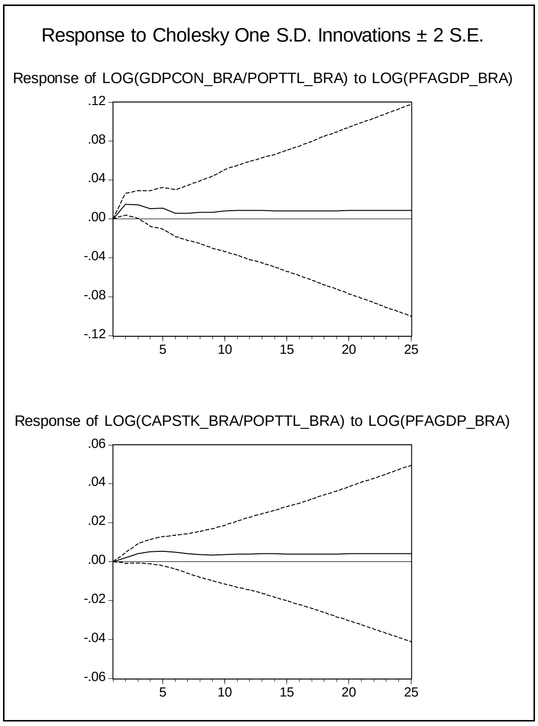

Figure 13. Impulse responses. Chile

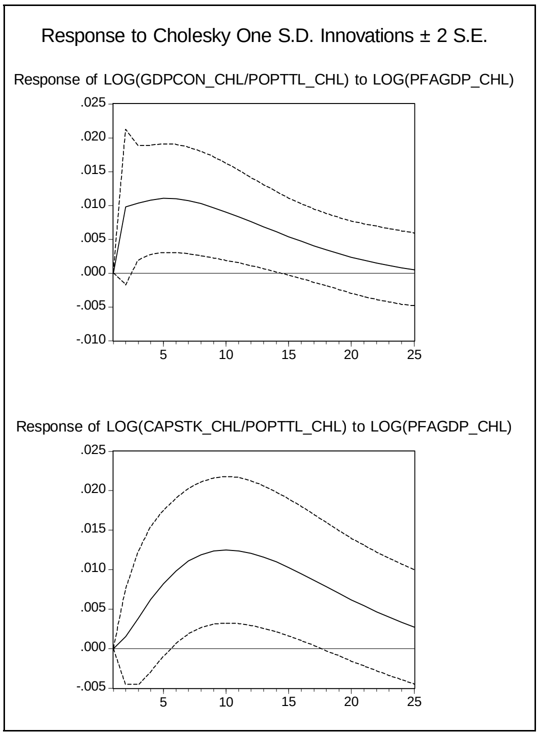

Figure 14. Impulse responses. Korea

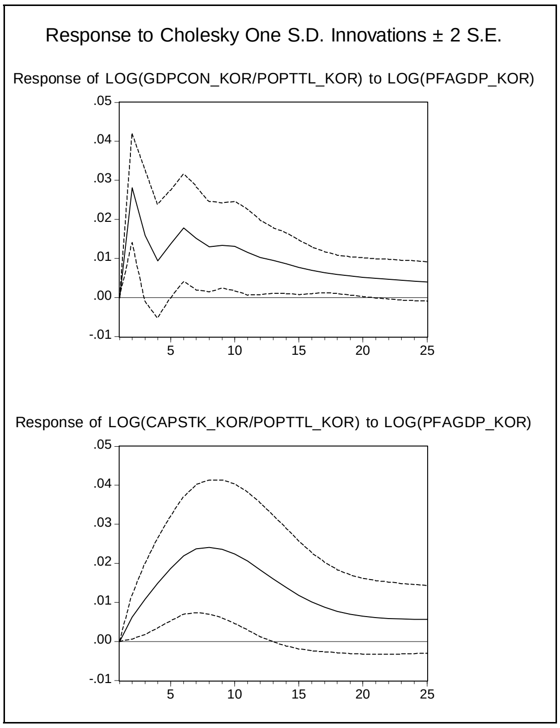

Figure 15. Impulse responses. Malaysia

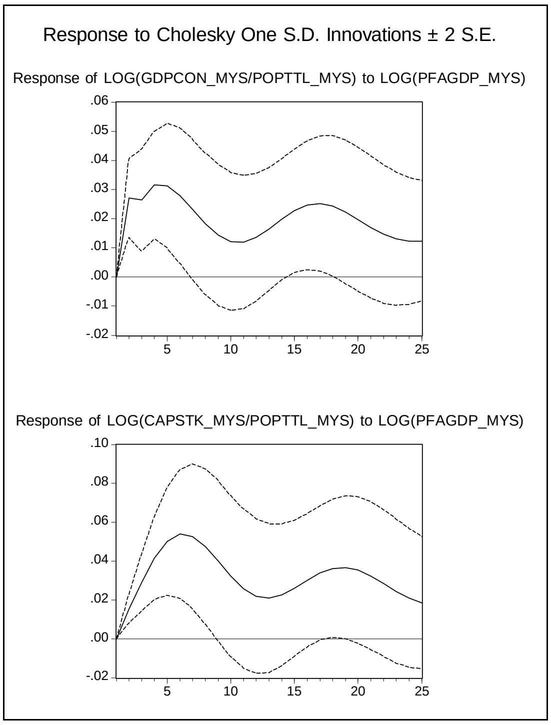

Figure 16. Impulse responses. South Africa