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Are funding of pensions and economic growth directly linked? New empirical results for some OECD countries

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ABSTRACT[1]

We empirically test on a panel of OECD countries the hypothesis of a direct and positive link between funding of pensions and economic growth, which is based on the idea that richer pension systems can accelerate the development of the financial system and thus promote a more efficient capital allocation. We follow Davis and Hu (2008) [Davis and Hu (2008), *Does funding of pensions stimulate economic growth?*, Journal of Pension Economic and Finance, Cambridge University Press, vol. 7 (02), 221-249] in estimating a modified Cobb-Douglas production function, where pension fund assets are treated as a shift factor, but we criticize their results from an econometric point of view, since both the Dynamic OLS and Mean Group (MG) estimators are inadequate in case of cross-sectionally correlated residuals. Indeed, we find a highly significant level of correlation in the MG residuals across countries

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that we attribute to common global shocks driving per capita outputs. Therefore we adopt a more general approach suitable to the presence of a multifactor error structure. Our results exclude the existence of a long run cointegration relationship between autonomous (or total) pension fund assets and per capita output for our panel of OECD countries, unless, in contrast to the conclusion of the cross-sectional dependence test, we ignore it and assume independence of residuals.

KEYWORDS: Pension funds assets, Output growth, Common factors, Heterogeneous panel, Panel cointegration, Panel spurious regression.

1. Introduction

The aging of the developed world's population has attracted the attention of Governments involved in social security policies. In fact it is widely accepted that, especially in developed countries, pay-as-you-go (PAYG) systems are no longer able to cope with demographic change. These systems are considered reasonably appropriate when economic growth and population growth are strong, but clearly this is not the case in many developed countries, for example those of continental Europe, where it is widely adopted (Boeri et al., 2006). Some important institutions, notably the World Bank (see Holzman and Hinz, 2005), sponsored a shift from PAYG system to the fully funded system, also for emerging economies. The Chilean pension reform of 1981 (see Holzman, 1997), that made the desired shift and that is considered by several as a decisive factor to make Chile the first South American member of the OECD, is famous in this respect.

Despite the arithmetic reasonableness of these considerations, and despite the shifting trend to funded systems which undoubtedly started around the world, there are still several countries that adopt the PAYG system. In fact, although there is a general consensus about the greater sustainability of the fully funded system with respect to the PAYG system, objections and reluctances lie mainly on the social costs of the transition. A typical remark is that the shift would create a “twice paying generation”, that would be required to pay for current pensioners, according to the current PAYG scheme, and also for funding their own pension plan. However, several criticisms have been made to this approach; see Feldstein and Samwick (1997). An interesting one focuses on the idea that a funded pension system would act as a stimulus to economic growth, and that this growth would thus be able to compensate for the losses incurred by the “twice paying generation”; see, e.g., Holzman (1997). Where does this belief originate?

Since Robert Solow (1956), economic theory has suggested a positive relationship between the saving rate of a country and its long-run output level. A correlation between the saving rate and the pension system has been widely considered and explored in the past decades, with mixed results depending on the methods used and the countries considered (see Kohl et al., 1998).

But there is also a more recent approach that has caught our attention. Some theories have hypothesized a *direct* link between the presence of pension funds in the economy and its growth rate. The origin of this link lies mainly on the idea that richer pension systems can accelerate the development of the financial system and thus promote a more efficient capital allocation.

A fully funded pension system, as mentioned above, would be more sustainable for public finances. This would result in a financial stability which, as the current European debt crisis revealed, is a very important factor for economic growth.

The establishment of a funded pension system leads to the presence of a large institutional investor in the market: the pension fund. The presence of institutional investors could, according to these theories, increase the demand for financial assets, both in the stock market and in the government bond market.

Moreover, compared to small investors, institutional investors could put pressure on those who organize and regulate the market, so that the presence of pension funds would better ensure the efficient functioning of markets.

Finally, through the participation share of their pension funds, reaching a partial mutual control, companies could achieve a broader and more loyal shareholder base. Also workers could be most interested in the performance of the company of which they are, indirectly, shareholders (see Blake, 1992). It is worth noticing that there are also some critical aspects concerning the presence of pension funds in the capital markets: they are mainly related to the degree of risk aversion of these institutions, for their social security role. The pension funds are heavily influenced by market trends (see for example Bikker et al., 2010), and this in turn could lead to pro-cyclical consequences, that could adversely affect subjects with a negative cycle.

In short, there is a theoretical justification, but subject to some caveats, for the existence of this direct link between the presence and value of pension funds in an economy and its growth rate and it seems to be reasonable. But what about empirical evidence?

Empirical works on this issue are not so numerous. An important contribution on this subject, besides that of Holzmann (1997), was made by E. Philip Davis (2004), through his studies on institutionalization. Especially interesting in this regard, however, is Davis and Hu (2008)'s paper. In this paper the authors specifically investigated the existence of the direct link between pension fund assets and economic growth through a modified Cobb-Douglas production function, considering pension fund assets as a shift factor. They found evidence of this relationship for both OECD countries and emerging economies using panel as well as country-by-country cointegration analysis. However, as they recognized, the relatively small number of observations casts some doubt on the robustness of their conclusions based on the dynamic heterogeneous panel model as well as on Johansen-cointegration tests. Moreover and most importantly, the Dynamic OLS results that Davis and Hu (2008) argue to be less affected by data limitations, are based on the critical assumption that the error terms are independently distributed across countries, so that the regression residuals shouldn't show any systematic pattern of correlation across countries. The problems arising from such correlation are well-known in the econometric literature on panel time series (Phillips and Sul, 2003; Andrews, 2005; Pesaran, 2006; Bai, 2009). Furthermore, in recent applied work it has been shown that cross-sectional correlation has a significant bearing on estimation (see e. g. Holly, Pesaran and Yamagata, 2010) and the results obtained in the empirical literature considering these issues have often eroded the significance of previous results (see for e.g. Eberhardt, Helmers and Strauss, 2013). For this reason and given the relevance of Davis and Hu (2008)'s result for its policy implications, it is interesting even with regard to this topic to revisit the previous empirical findings in a more modern and robust econometric perspective. In particular, we investigate the adequacy of the implicit assumption on which their panel data analyses were based, taking into account the alternative hypothesis of error cross-sectional dependence.

The latter would suggest the presence of common latent factors, that affect all countries albeit to a different extent. As emphasized above, testing for cross-sectional independence is crucial for the validity of the results obtained by Davis and Hu (2008), since both the DOLS and Mean Group estimator they applied turn out to be inconsistent under the alternative hypothesis. We find a highly significant level of correlation in the residuals across countries. Therefore, in order to account for it we estimate the long-run relationship between pension funds and output per capita, considered by Davis and Hu (2008), in the presence of a multifactor error structure. We use the CCEMG and CCEP estimators advanced by Pesaran (2006) and Kapetanios, Pesaran and Yamagata (2011). Moreover, we test for the possibility of a panel spurious regression using the procedure reported in Holly, Pesaran and Yamagata (2010) that, more recently, Banerjee and Carrion-i-Silvestre (2011) have shown to be consistent also under the null hypothesis of a unit root in the idiosyncratic errors.

Lately, Zandberg and Spierdijk (2013) have criticized the conclusions drawn by Davis and Hu (2008), but on different grounds. One of their criticisms concerns the formulation of the pension assets variable. In fact, Davis and Hu (2008) considered only the autonomous pension fund assets, while Zandberg and Spierdijk (2013) consider as more appropriate variable the total pension fund assets[2]. To verify possible differences in the results, we will also take account of Zandberg and Spierdijk (2013)'s suggestion reestimating the model with their variable.

2. Model specification and data description

As mentioned in the previous section, our empirical analysis is based on the model specification adopted by Davis and Hu (2008). They considered the following standard Cobb-Douglas production function, normalized by labor force, with the addition of the pension fund assets as a shift factor:

$$Q_{i,t} = e^{\alpha_i+\gamma_i t+e_{i,t}} x K_{i,t}{}^{\beta_{i1}} x P_{i,t}{}^{\beta_{i,2}} \tag{1}$$

where Q is output per unit of labor, K is capital per unit of labor and P denotes the pension fund assets. Expressing the model in log terms we obtain:

$$lQ_{i,t} = \alpha_i + \gamma_i t + \beta_{i1} lK_{i,t} + \beta_{i2} lP_{i,t} + e_{i,t} \tag{2}$$

Where $\alpha_i+\gamma_i t$ represents the technology level with α_i the individual intercept term and t the time trend, and e_{it} is an error term.

As noted above, Davis and Hu (2008) estimated the heterogeneous long-run relationship in eq. (2) by employing the dynamic panel data model proposed by Pesaran and Smith (1995) and their Mean-Group (MG) estimator, under the assumption that the long-run elasticities were random coefficients with common mean and that the error terms were independently distributed across countries. It is worth noticing that the assumption of cross-sectional independence of the error terms is critical for the consistency of the average long-run estimates obtained by the MG estimator (see Pesaran, 2006, Kapetanios, Pesaran and Yamagata, 2011). So, it is of vital importance to test for residual cross-sectional independence if misleading conclusions are to be avoided. In the following we will analyze this critical issue in more depth.

[2] When we refer to P we consider pension assets of autonomous pension funds, when we refer to TP we consider also pension assets of funds managed by other institutions such as insurance companies, banks, investment companies, etc.

Moreover, before estimating the long-run relationship of interest, Davis and Hu (2008) performed three types of panel unit root test on the variables involved: the LLC test (Levin et al., 2002), the IPS test (Im et al., 2003) and the Hadri (2000)'s panel stationarity test. Also in this case, the adoption of the above tests implicitly suggests the assumption of errors cross-sectional independence, because otherwise their results would be misleading.

Therefore, in the next section we firstly apply the IPS panel unit root test to *lK*, *lP* and *lQ* and then check for residuals cross-sectional dependence using the Pesaran (2004)'s CD test.

Unfortunately, we do not have access to the data-set used by Davis and Hu (2008), but we use the same sources as regards both the pension data and output per capita. We analyze 16 OECD countries, listed in Appendix A together with the time span of the variables. Pension data are taken from OECD's *Institutional Investors* database. Capital stock data are taken from the Annual Macro-economic Database of the European Commission's Directorate General for Economic and Financial Affairs. Data on per capita output come from the World Development Indicators of the World Bank, whereas data on total pension assets are provided by Zandberg and Spierdijk (2013), and were taken from several OECD collections. In this case, due to a different time coverage of the variable, the sample is restricted to 12 OECD countries.

3. Unit root and cross-sectional dependence tests

We start the empirical analysis applying the IPS unit root test to the three variables under scrutiny[3]. The results are reported in Table 1. For the variables in levels, the IPS test is performed with a constant and a linear trend as deterministic components, whereas only a constant term is added to the equation specification for the first differences of the variables. Moreover, the test statistics are reported for both models with one or two lags of the dependent variable as regressors.

[3] Like Davis and Hu (2008), per capita output and capital are used as proxies for respectively output and capital per units of labor. The ratio of autonomous pension fund assets to GDP proxies the value of the pension fund assets in the economy.

TABLE 1. IPS unit root test for the three variables of the model

Im, Pesaran and Shin W-stat				
Intercept and trend				
Lags	1	P-value	2	P-value
lP	0.9309	0.8244	0.64544	0.7407
lK	-0.39131	0.3478	1.94065	0.9738
lQ	-0.4953	0.3102	1.04212	0.8513
Intercept only				
ΔlP	4.54761	0	3.86506	0.0001
ΔlK	-2.91882	0.0018	-2.08863	0.0184
ΔlQ	-6.25772	0	-3.47467	0.0003

As we can see from Table 1, all the three variables appear to be I(1). However, the results of the CD test reported in Table 2 strongly reject the null hypothesis of cross-sectional independence of the residuals, casting doubt on the reliability of the conclusions reached by the IPS test.

TABLE 2. Pesaran's CD test for the single variables of the model

Pesaran's CD test		
Intercept and trend	p-value	p-value
Lags	1	2
lP	0.0048	0.0008
lk	0.0007	0.0001
lq	0.0023	0.0001

Therefore, in Table 3 we report the results of the CIPS test which Pesaran (2007) has shown to be robust to the presence of cross-sectionally correlated errors. In order to take account of error cross-sectional dependence, as possibly deriving from either an omitted factor structure or from spatial spillovers, Pesaran (2007) proposed to augment the regressors of the IPS test with cross section averages of both the regressors and the dependent variable. From Table 3 we can see that the null hypothesis of a unit root is not rejected for all the three variables also in this case.

TABLE 3. CIPS unit root test for the three variables of the model

Pesaran's CIPS test		
Intercept and trend		
Lags	1	2
lP	-2.37	-1.21
lk	-1.47	-0.76
lq	-1.76	-0.91
Critical values (5%)	-2.76	-2.76
Critical values (10%)	-2.66	-2.66
Intercept only		
Lags	1	2
ΔlP	-2.34**	-1.55
Δlk	-1.98	-0.99
Δlq	-3.12**	-1.85
Critical values (5%)	-2.25	-2.25
Critical values (10%)	-2.14	-2.14

** stands for significant at 5%.

The CIPS test results for the differenced variables are instead mixed. The presence of a unit root is rejected for ΔlP and ΔlQ but only for the model specification with one lag. However, this might depend on a loss of power when adding further lags: given the short sample and the fact that we are testing for a unit root the variables in first differences, we consider the test results with one single lag the most reliable. As far as *lK* is concerned, CIPS tests conclude for the presence of at least two unit roots, which both for theoretical considerations and on inspection of the variable graphs (see Figure 1) seems to be unrealistic. Therefore, all three variables will be treated as I(1) in the subsequent analyses.

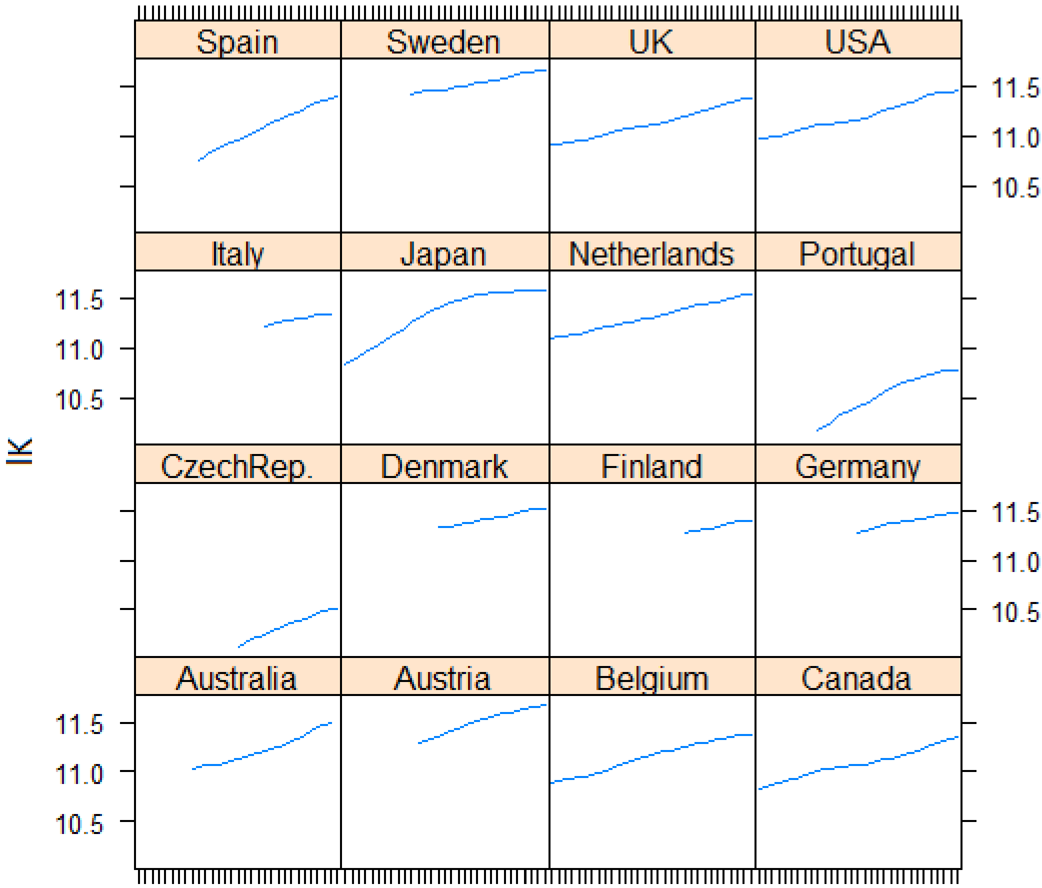

FIG. 1. Log of real capital per capita for 16 OECD countries, 1980-2010

4. Estimation and cointegration analysis

In Table 4 we report the MG estimates of the long-run elasticity obtained by directly estimating eq. (2) separately for each country by OLS and averaging the coefficients estimates across countries. Notice that both estimated elasticities are positive but only the capital elasticity is significantly different from zero. However, as noted by Pesaran (2006) and Kapetanios, Pesaran and Yamagata (2011), the assumption of cross-sectional independence of the error terms in eq. (2) is critical for the consistency of the average long-run coefficient estimates obtained by the MG estimator[4].

[4] However, the MG estimator turns out to be consistent in case of residuals autocorrelation and/or heteroskedasticity.

TABLE 4. Estimates of the model (s.e in parentheses)

	$\hat{\alpha}$	$\widehat{\beta_1}$	$\widehat{\beta_2}$
Intercept and trend			
Mean Group	-8.11 (4.79)	1.61 (0.42)	0.009 (0.027)
CCEMG	5.29 (3.86)	0.62 (0.30)	0.03 (0.018)
CCEP		0.38 (0.32)	0.06 (0.041)
Intercept only			
Mean Group	-2.14 (1.06)	1.07 (0.10)	0.009 (0.027)
CCEMG	0.38 (1.19)	1.09 (0.23)	0.076 (0.021)
CCEP		0.73 (0.59)	0.06 (0.062)

Therefore in Table 5 we report the results of the CD test applied to the OLS residuals of eq. (2) estimated with and without the linear trend.

TABLE 5. Pesaran's CD test for the OLS residuals of the model with trend and/or intercept

Pesaran's CD test		
	CD Statistic	P-value
Intercept and trend	5.23	0.000
Intercept only	13.38	0.000

As can be seen from Table 5 we are in presence of cross-sectional dependence in both cases. A result this one that suggests the existence of common latent factors which drive the per capita outputs of the OECD countries under analysis.

To take account of error cross-sectional dependence, Pesaran (2006), Kapetanios, Pesaran and Yamagata (2011) and Pesaran and Tosetti (2011), among others, have proposed the following multifactor error structure for the disturbances of eq. (2):

(3) $$e_{it} = \lambda_i^{'} f_t + \varepsilon_{it}$$

where f_t is the vector of unobserved common factors and λ_i represents its corresponding vector of factor loadings. The factor loadings are assumed to be heterogeneous across countries which means that each single common factor may have a different impact on the OECD's per capita outputs. The remainder idiosyncratic error, ε_{it}, is allowed to be a general stationary process[5], as well as being weakly cross-sectionally dependent (see Pesaran and Tosetti, 2011). Notice that in eq. (3) the common factors take account of the strong (i.e. global) forms of cross-sectional dependence, whereas the idiosyncratic errors take account of the residual weak (i.e. local) forms of dependence across countries.

Moreover, the unobserved common factors could be correlated with the regressors of eq. (2) and, given the results of the panel unit root tests reported above, they may be also I(1). Therefore, the following general specification is proposed for the regressors:

$$x_{it} = a_i + A_i f_t + u_{it} \quad (4)$$

where $x_{it} = \left(lK_{it}, lP_{it}\right)'$, a_i and A_i are a $2x1$ vector of individual specific intercepts and the factor loading matrix, respectively. Finally, the vector of disturbances, u_{it}, is assumed to be a general stationary process.

Kapetanios, Pesaran and Yamagata (2011) have shown that the common correlated effects (CCE) estimators proposed by Pesaran (2006) are still valid[6] when the unobserved common factors contain unit root processes. It is worth noticing that in such case y_{it}, x_{it} and f_t must be cointegrated. Moreover, they have shown that both the Common Correlated Mean Group (CCEMG) and the Common Correlated Effects Pooled (CCEP) estimators turn out to be consistent and asymptotically Normally distributed[7] for the mean of the individual specific slope coefficients in eq. (2).

Therefore we proceed by calculating the CCEMG and CCEP estimates of the model. The estimates are shown in the Table 4.

We notice that the calculated elasticities vary widely considering the two estimators, but only the CCEMG estimate of the mean of β_{2i}, named β_2, for the model without trend appears to be significantly different from zero as well as positive. However, given that the dependent variable was found to be I(1), we tend to prefer the estimation results of the model without deterministic trends. For this model we find that also the average long-run capital elasticity is positive and significantly different from zero when the slope coefficients are assumed to be heterogeneous but random across countries, instead of homogeneous as assumed by the CCEP estimator.

In order to avoid estimating a panel spurious regression[8], it is crucial also in this context to test for cointegration. Therefore in Table 6 we report the results of the CIPS test for the presence of a unit root in the idiosyncratic residuals[9], $\hat{\varepsilon}_{it}$. In addition to the standard

[5] So that in general it will be autocorrelated.
[6] Pesaran (2006) noted that linear combinations of the unobserved factors can be well approximated by cross-section averages of the dependent and the observed regressors and proposed a new set of estimators, referred to as CCE estimators which are computed by running standard panel regressions augmented with these cross-section averages.
[7] Under the assumption of random slope coefficients.
[8] Panel spurious regression has been tackled in Phillips and Moon (1999).
[9] Therefore the variables result to be not cointegrated under the null hypothesis of a unit root in the residuals.

procedure, the CIPS test is also performed using the test procedure proposed by Banerjee and Carrion-i-Silvestre (2011) and their critical values[10].

TABLE 6. CIPS test for the CCEMG and CCEP residuals of the model

Pesaran's CIPS test		
Intercept only		
Lags	1	2
CCEMG residuals	-0,61	-0,44
CCEP residuals	-1,21	-1,69
Critical values (5%)	-2,25	-2,25
Critical values (10%)	-2,14	-2,14
BC test statistic	-0,21	0,06
Critical values (5%)	-2,32	-2,32
Critical values (10%)	-2,22	-2,22
Intercept and trend		
Lags	1	2
CCEMG residuals	-0,83	-1,18
CCEP residuals	-0,54	-0,29
Critical values (5%)	-2,76	-2,76
Critical values (10%)	-2,66	-2,66
BC test statistic	-0,23	-0,09
Critical values (5%)	-2,91	-2,91
Critical values (10%)	-2,82	-2,82

Notice that the IPS test would be inconsistent[11], according to the CD test results reported in Table 7, when applied to the model residuals computed gross of potential common factors.

[10] Such a procedure generalizes that used by Holly et al. (2010) for testing for panel cointegration. More importantly, Banerjee and Carrion-i-Silvestre (2011) show that the CCEP estimator is still consistent for the long-run average coefficient regression in the presence of a panel spurious regression.
[11] As such, it might lead to misleading conclusions about the existence of a cointegration relationship.

TABLE 7. CD test on the model residuals (gross of potential common factors)

	CD statistic	P-value
CCEMG intercept only	15,4811	0.00
CCEMG intercept and trend	18,4492	0.00
CCEP intercept only	8,1021	0.00
CCEP intercept and trend	27,0135	0.00

From Table 6 we can see that we never reject the null hypothesis of no cointegration.

5. Results for the total pension fund assets

As noted in the Introduction, also Zandberg and Spierdijk (2013) have criticized the conclusions drawn by Davis and Hu (2008), but on different grounds. One of their criticisms concerns the formulation of the pension assets variable. In fact, Davis and Hu (2008) considered only the autonomous pension fund assets, while Zandberg and Spierdijk (2013) consider the total pension fund assets (TP) as more suitable for the analysis, since they note that the beneficial effect of the investment exists also if it's not a pension fund but another type of institution that carry out it. Therefore, we check the robustness of the conclusions reached till now by estimating the model above with the variable proposed in Zandberg and Spierdijk (2013). Unfortunately, as anticipated above, the sample has to be restricted to 12 OECD countries, due to the different time coverage for this variable.

The estimates are shown in Table 8.

TABLE 8. Estimates of the model with total pension assets (*TP*) as regressor.

	$\hat{\alpha}$	$\widehat{\beta_1}$	$\widehat{\beta_2}$	IPS/CIPS statistic for the residuals of the model
Intercept and trend				
Mean Group	-26.09 (8.47)	3.22 (0.75)	0.08 (0.05)	3,355***
CCEMG	-1.55 (6.17)	0.90 (0.93)	0.006 (0.15)	-0.67
CCEP		-0.23 (0.97)	0.007 (0.015)	-2.81
Intercept only				
Mean Group	0.78 (2.60)	0.81 (0.22)	0.005 (0.04)	-1,875*
CCEMG	-2.37 (2.86)	0.31 (0.70)	0.02 (0.02)	-1.39
CCEP		-0.20 (0.58)	0.001 (0.015)	-3.24

*** stands for significant at 1%, * at 10%

We emphasize in particular that the estimates of the *lTP* long run coefficient are not significantly different from zero. Moreover, IPS and CIPS unit root tests identify a cointegration relationship only with regard to the MG estimates. But these estimates are not reliable, because the CD test revealed the presence of cross-sectional correlation, as you can see in table 9.

TABLE 9. Pesaran's CD test for the residuals of the model with trend and/or intercept

Pesaran's CD test		
	CD Statistic	P-value
OLS Intercept and trend	5,23	0,00
OLS Intercept only	13,38	0,00
CCEMG intercept only	15,4811	0,00
CCEMG intercept and trend	18,4492	0,00
CCEP intercept only	8,1021	0,00
CCEP intercept and trend	27,0135	0,00

So, even in this case, when properly accounting for cross-sectional correlation we do not find evidence of a long-term relationship between pension assets value and economic growth. This leads us to stress once again the importance of considering cross-sectional correlation of errors when it is present. Secondly we conclude that, although the observations of Zandberg and Spierdijk (2013) about the specification of the pension variable seem well-founded, the results obtained taking them into account lead us to the same conclusions. However, it is important to stress that the brevity of the sample makes it very difficult to draw definite responses.

6. Conclusions

The significance of the previous empirical results obtained in the literature, using panel time series models under the assumption of errors cross-sectional independence, has been often eroded by subsequent analyses, where the presence of cross-sectional correlation was tested and properly accounted for. A similar situation has been encountered in our context, since we have found that cointegration between pension fund assets and per capita output has not been rejected for our panel of OECD countries, but only under the false hypothesis of independence of residuals across countries. Therefore, taking into account errors cross-sectional dependence we excluded the existence of a positive and direct relationship between the value of pension fund assets and output, at least regarding OECD countries.

We cannot exclude the possibility that the negative effects on growth of pension funds' risk aversion offset the positive ones discussed by Davis and Hu (2008), and reviewed in the Introduction. In fact, also some advocates of the positive effect on growth of pension funds recommend a different way of allocating investments (see Hu, 2006).

However, the small sample size issue due to the scarcity of data has to be emphasized. In fact, the analyses we carried out were based on an unbalanced panel, where – in the case of autonomous fund assets – only ten out of sixteen developed countries cover a full thirty-years time period.

In the other case, when we considered total pension assets, we had a balanced panel of twelve countries, only ten-years long.

We believe that a longer sample is a necessary condition in order to draw more significant and definitive conclusions on the long run relationship between pension fund assets and growth. Therefore, although the results we obtained with the data currently available have not found evidence of such a relationship, we hope that in the coming years a greater amount of pension data will be recorded, in order to get more comprehensive answers to the questions we posed.

APPENDIX A

DATABASE 1

Country	Time period	
Australia	1988	2009
Austria	1991	2010
Belgium	1980	2010
Canada	1980	2010
Czech Republic	1995	2010
Denmark	1994	2010
Finland	2000	2010
Germany	1995	2010
Italy	1999	2009
Japan	1980	2010
Netherlands	1980	2010
Portugal	1989	2010
Spain	1989	2010
Sweden	1990	2010
United Kingdom	1980	2010
United States	1980	2010

DATABASE 2 (i.e. "Total pension assets" data)

Country	Time period	
Austria	2001	2010
Belgium	2001	2010
Canada	2001	2010
Czech Republic	2001	2010
Denmark	2001	2010
Finland	2001	2010
Germany	2001	2010
Netherlands	2001	2010
Portugal	2001	2010
Spain	2001	2010
Sweden	2001	2010
United States	2001	2010

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