

Österholm, Pär

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Department of Economics

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Pär Österholm

Department of Economics
Uppsala University
P.O. Box 513
SE-751 20 Uppsala
Sweden
Fax: +46 18 471 14 78

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PÄR ÖSTERHOLM

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Estimating the Relationship between Age Structure and GDP in the OECD Using Panel Cointegration Methods*

Pär Österholm*

September 2004

Abstract

Economic theory suggests that variations in countries' age structure should affect the economy on an aggregate level. This paper investigates the relationship between age structure and GDP in 20 OECD countries using annual data from 1970 to 1999. Using new methodology, the relationship between the variables can be formulated in levels despite the presence of unit roots in the time series. Applying two panel cointegration tests proposed by Pedroni (1995, 1997a, 1999), support is found for a long run relationship between GDP and the number of people in five different age groups. Coefficient estimates from panel regressions support effects in line with the life cycle hypothesis and human capital theory; children and retirees are found to have a negative or relatively smaller positive effect on GDP than productive age groups.

JEL Classification: C23, J10

Keywords: Age structure, GDP, Panel cointegration

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* Department of Economics, Uppsala University, Box 513, S-751 20 Uppsala, Sweden
e-mail: par.osterholm@nek.uu.se

1 Introduction

If the life cycle hypothesis of Modigliani and Brumberg (1954) and the human capital theories of Schultz (1960) and Becker (1964) are to be believed, the economic behaviour of an individual will change with age. For instance, these theories imply that in some stages of life, the individual will be a net borrower whilst at others a net saver. In addition, to the extent that education and experience vary with age, so to will the productivity of a worker. Whilst these theories yield interesting microeconomic predictions, their relevance is not confined to the individual. Aggregating over all individuals in an economy, and adding the empirical phenomenon of varying cohort sizes in the population, these theories make important predictions about relationships between the age structure of the population and a number of macroeconomic variables.

In earlier research that applied the predictions of the life cycle hypothesis and human capital theories to macroeconomics, the first variable to receive attention was aggregate saving in which Leff (1969), Bentzel and Berg (1983) and Horioka (1989) found age structure effects. Aggregate saving seems an obvious starting point for this field of research given its direct relation to the life cycle hypothesis, but later research has widened its scope. Age structure effects have been established on consumption by Fair and Dominguez (1991), on real interest rates by McMillan and Baesel (1990) and Lenehan (1996) and on the current account by Taylor and Williamson (1994) and Herbertsson and Zoega (1999).

Another macroeconomic variable that has received a fair amount of attention in the macroeconomics and demography literature is the growth of GDP. McMillan and Baesel (1990), Lenehan (1996) and Bloom and Sachs (1998) are just a few who have found significant age structure effects on growth. Previous studies have established that the dependency ratio of a country affects its growth rate, and whilst it seems intuitive that a country with a larger fraction of children or retirees should exhibit lower growth than a country whose population consists mainly of people in productive ages, a more straightforward approach is to study directly the relationship between the level of GDP and age structure measured in number of people. One obvious reason why this method

has until now not been considered is that the methodology to deal with nonstationarity of series in levels – especially in a panel setting – has been developed only recently.

This paper aims to further investigate the impact of age structure on the economy by estimating the relationship between population and GDP using a panel of 20 OECD countries. The econometric specification, in which GDP is explained by population distributed amongst five different age groups, is interpreted as a reduced form but could also be consistent with an aggregate production function under some assumptions. Given the likely presence of unit roots in both GDP and population, two panel cointegration tests developed by Pedroni (1995, 1997a, 1999) are used to establish whether the variables are cointegrated and the parameters of the model are then estimated using pooled OLS and fixed effects following Phillips and Moon (1999, 2000).

Estimation of reduced form specifications is common in the empirical literature when demographic effects on macroeconomic variables are investigated and has potential advantages. Unlike many macroeconomic variables that are generally difficult to forecast at longer horizons, the demographic structure can be forecasted at long range with high precision. Many reduced form specifications – including the one in this paper – have no other explanatory variables than population; given that there are stable correlations between age structure and the macroeconomic variable at hand, reduced form models and demographic data could then serve as useful forecasting devices for the medium- and long-run trends in economic variables as argued in Lindh and Malmberg (2000), Andersson and Österholm (2001) and Lindh (2004). A relationship between GDP and age structure is therefore interesting not only to test the relevance of life cycle and human capital effects on an aggregate level; it could also be a tool that improves forecasts of GDP.

This paper is organised as follows. Section 2 discusses the theoretical foundations behind the model and Section 3 addresses econometric issues. In Section 4 data are described and analysed, cointegration tests are performed, the model is estimated and the results are discussed. Finally, Section 5 concludes.

2 Age structure effects on GDP

The traditional way of modelling demography in macroeconomics, such as in Ramsey (1928), Solow (1956) and Romer (1990), is to assume that population growth is stable and given by an exogenous exponential rate, n . Though convenient, and perhaps a good approximation in the very long run, this is not realistic for shorter horizons, as evidenced by the well established fact that fertility is time varying. This variation leads to periods of baby booms and baby busts, which can clearly be seen in the variations in cohort size depicted in Figures 1 to 3 below. Given the very large differences in cohort sizes shown, it comes as no surprise to economists and policymakers alike that the age structure can affect the economy.

The common usage of reduced forms when studying the effects of age structure on macroeconomic variables is convenient because it can incorporate both direct and indirect effects. This paper takes as its starting point equation (1) below,

$$y_{i,t} = \delta_i + \phi_i t + \gamma_1 l_{0-14i,t} + \gamma_2 l_{15-29i,t} + \gamma_3 l_{30-49i,t} + \gamma_4 l_{50-64i,t} + \gamma_5 l_{65-i,t} + \varepsilon_{i,t}$$
$$i = 1, \dots, N \text{ and } t = 1, \dots, T \quad (1)$$

where $y_{i,t}$ is the logarithm of GDP for country i at time t , $l_{aa-bbi,t}$ is the logarithm of the number of people between age aa and bb in country i at time t , N is the size of the panel and T is the number of observations over time.

Figure 1. Population age structure of the United States in 2000. (Population in thousands).

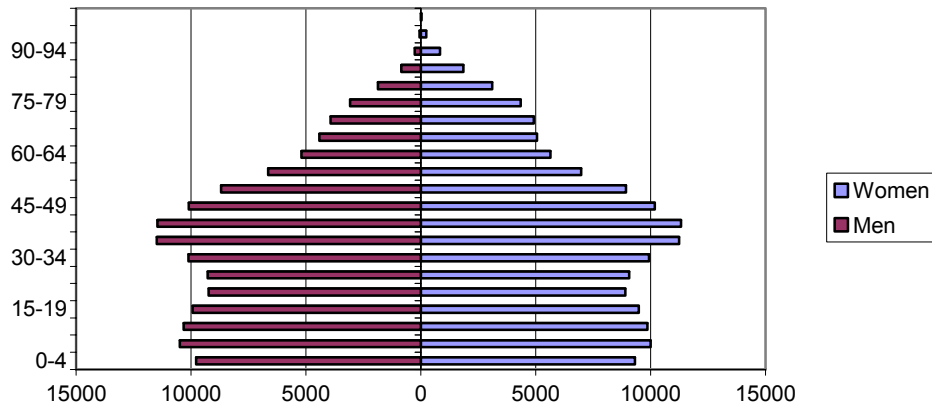


Figure 2. Population age structure of Japan in 2000. (Population in thousands).

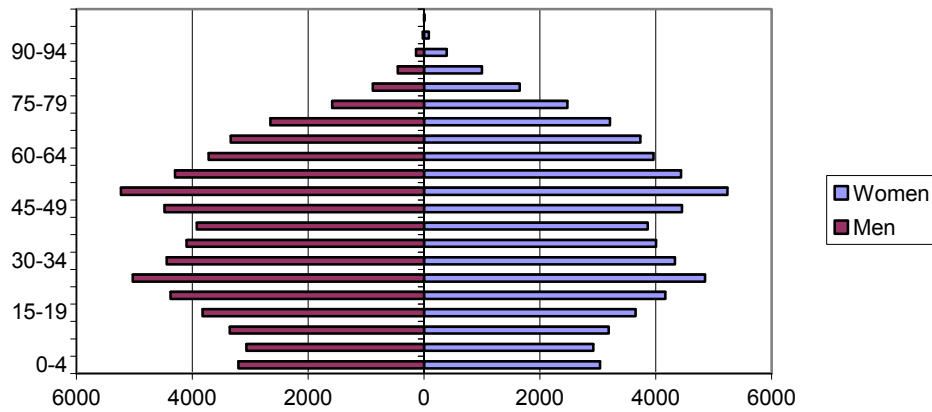
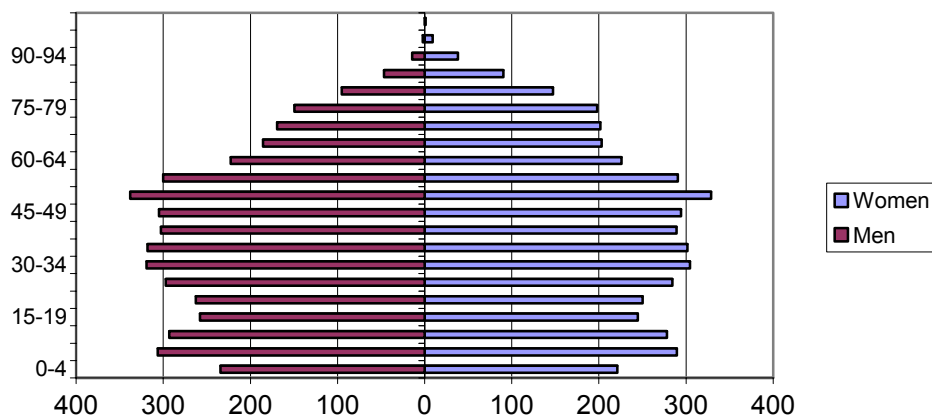


Figure 3. Population age structure of Sweden in 2000. (Population in thousands).



The econometric details of this specification will be discussed in Sections 3.3 and 3.4 whereas here we will discuss the kind of effects we expect to catch with this formulation and under which circumstances this approximation is reasonable. Regarding equation (1) as a reduced form of the structural relationships in a macroeconomy, it can reasonably be used to estimate the effects that age structure has on a linearised version of the economy. In doing so, we abstract from a well specified model and are mainly concerned with whether age structure affects GDP at all and the approximate size of those effects. Our aim here is not to address the underlying structural mechanisms.

What effects can then be expected? Using the national income identity $Y = C + I + G + NX$, GDP can be decomposed into four categories, each of which has been shown to be affected by demography. Previous research has to a large extent been supportive of the implications of the life cycle hypothesis. As expected, countries with a high proportion of very young or old people have been found to have low or negative saving ratios and current accounts since these groups must borrow or dissave to smooth consumption (Taylor and Williamson, 1994, Herbertsson and Zoega, 1999). Furthermore, children and retirees tend to consume more public goods and services and do not, in general, contribute greatly to production. In line with this, it has been argued that children and retirees are a source of inflationary pressure (McMillan and Baesel, 1990, Lenehan, 1996, Lindh, 2004). Higher real interest rates (McMillan and Baesel, 1990, Lenehan, 1996) and lower investment (Lindh and Malmberg, 2003) are also observed in economies with more children and retirees, which is reasonable considering the likely unwillingness of retirees to postpone consumption.

Qualitatively summing up these age structure effects on macroeconomic variables, we would expect children and retirees to have negative or slightly positive effects and young adults and middle aged to have more positive effects on GDP. However, counteracting effects might be found in G , since more young and/or old implies higher government expenditures. This has the potential to increase GDP, but the outcome is not obvious. For example, crowding out could exert an offsetting effect on investment.

It is useful to consider how the reduced form described above can be understood within the framework of an aggregate production function. This helps to clarify some details of the specification and offers a structured way of explaining potential problems. From a production function point of view, the link between population and output is direct, given that labour is a vital input in most production functions. The simplest case for which equation (1) could be valid is when GDP is produced by a Cobb-Douglas type production function with only labour input,

$$Y_{i,t} = A_i L_{i,t}^\gamma \quad (2)$$

where A_i is a scaling parameter describing technological efficiency and $L_{i,t}$ is labour in country i at time t . Taking logarithms of (2) and adding an error term to the model, the correct econometric specification would then become

$$y_{i,t} = a_i + \gamma l_{i,t} + e_{i,t} \quad (3)$$

where lower-case letters represent the logarithm of upper-case letters. With homogeneous labour in the production function, it is clear that dividing the labour inputs in equation (1) into age categories is unnecessary. Not only will the coefficients on children and retired be zero since they are not part of the labour force, but we will also find that $\gamma_2 = \gamma_3 = \gamma_4$. Though equation (1) is consistent with such a production function, in the case of homogeneous labour it has redundant regressors – including the time trend – which reduces the precision of the estimates. The age-specific effects posited by equation (1) are therefore clearly consistent with more complex production functions that allow for different types of labour, such as that suggested in equation (4).

$$Y_{i,t} = A_i L_{15-29i,t}^{\gamma_2} L_{30-49i,t}^{\gamma_3} L_{50-64i,t}^{\gamma_4} \quad (4)$$

According to human capital theory we should expect different age groups to exhibit differing degrees of productivity; both schooling and experience are generally

considered to increase productivity.¹ Following Griliches and Mairesse (1995), who point out that it is important to allow for differences in the quality of labour when estimating production functions, in this paper quality differentials are proxied by the productivity differences assumed to characterise different age groups in the population.

To this point, capital has been a striking omission from the posited production function. Once capital is introduced into the model, the standard approach is to employ a Cobb-Douglas function such as that given in equation (5).

$$Y_{i,t} = A_i L_{i,t}^\gamma K_{i,t}^\beta \quad (5)$$

where $K_{i,t}$ is physical capital. Rewriting into logarithms and adding an error term yields

$$y_{i,t} = a_i + \gamma l_{i,t} + \beta k_{i,t} + e_{i,t} \quad (6)$$

which, given data on labour and capital, appears to be a straight forward equation to estimate.² This specification has been used both at macro- and microeconomic levels. One of the main reasons for its popularity has been plausible parameter estimates in line with both constant returns to scale and evidence from factor shares.³

Comparing equations (3) and (6), the ramifications of misspecification are transparent. If data are generated according to equation (6) but equation (3) is estimated, the estimator of γ is both biased and inconsistent. It can be shown that unless capital is uncorrelated with labour – which most economists would agree is unlikely – the effect of capital on GDP will be included in the coefficient on labour. This is one reason to avoid giving structural interpretations to parameters in a reduced form estimation such as equation (1).

¹ The effect of experience is generally modelled as a concave function where the worker's stock of human capital peaks in the middle ages, towards the end of the working years.

² It should be pointed out that because of the likely possibility of endogeneity of regressors, OLS will generally not yield consistent estimates.

³ Work taking its starting point as the Cobb-Douglas production function includes, for instance, Blundell and Bond (2000).

The problems with omitted variables have so far been discussed under the assumption that all variables are stationary. If variables are in fact generated by unit root processes, the omission of a relevant explanatory variable becomes a more serious issue. In a single equation time series framework, not only will the problems of bias and inconsistency remain, but the regression has a higher chance of being spurious with the attendant consequences of diverging t - and F -statistics.⁴ There are some circumstances, however, in which the omission of relevant explanatory variables need not be particularly severe. Assume that production of GDP can be represented by a Leontief production function as given by equation (7).

$$Y_{i,t} = A_i \min(dL_{i,t}, cK_{i,t}) \quad (7)$$

Since capital and labour always are used in fixed proportions in this economy, there is no loss in fit or predictive power from excluding capital from an estimated model. Nor will the regression be spurious in the case when integrated time series data are present. The inconsistency of parameter estimates, however, is still an issue and the adequacy of the formulation then depends on the purpose of the estimation.

The same argument applies to all homothetic production functions under the assumption of constant relative prices and could justify the usage of equation (1) even if the production function is of Cobb-Douglas type with both varying qualities of labour and one type of capital as inputs.⁵ Given that labour and capital are used in fixed proportions, the omission of capital might actually improve the model, for in the first place, it is not clear which measure of capital should be used in the estimation (Menderhausen, 1938) and secondly, the capital stock is notoriously difficult to measure accurately.⁶

The technological efficiency parameter A_i also deserves attention. A large literature has been devoted to modelling technological change, and as Griliches and Mairesse (1995)

⁴ See for instance Phillips (1986).

⁵ This might be a brave assumption to rest upon though; there is strong evidence that real wages are trending upwards over time whereas real interest rates are not.

⁶ See for instance Hulten and Wykoff (1996).

pointed out, technical change should not be ignored when estimating production functions. When A_i is assumed to be time-dependent, the above equations (2), (4), (5) and (7) can be written succinctly as a production function in the following form

$$Y_{i,t} = A_{i,t} f(\cdot) \quad (8)$$

where $A_{i,t}$ is an index of technological efficiency in country i at time t . Adopting the common assumption of a log-linear exogenous trend in technological efficiency – consistent with most of the exogenous growth literature as employed by Ramsey (1928) and Solow (1956) – together with a production function of Cobb-Douglas type with only homogeneous labour, the appropriate econometric specification becomes⁷

$$y_{i,t} = c_i + b_i t + \gamma_{i,t} + e_{i,t}. \quad (9)$$

In this situation where technology grows at a constant, exogenous rate, a time trend is sufficient to capture technological change and the econometric specification in equation (1) is consistent with the theoretical relationship in equation (9), albeit allowing for more diversity in labour inputs than (9) suggests.

Overall, the econometric specification of equation (1) can be considered a fair approximation to a number of underlying production functions. Omission of relevant variables in the estimation, particularly capital, may seem like a potential obstacle but this variable in particular is likely to be closely related to the age structure and its effects captured by already included variables. Not only have age structure effects been established on investments empirically, it is also clear that capital and labour are used in fixed proportions under certain conditions presented above.⁸ The possible omission of relevant explanatory variables is another rationale for including groups that are not in

⁷ In a discrete time set up where $(A_t - A_{t-1})/A_{t-1} = g$, the log-level of A_t can be expressed as $a_t = a_0 + t \ln(1 + g)$ using recursive elimination. In continuous time where $\dot{A}_t/A_t = g$, the corresponding expression is $a_t = a_0 + tg$.

⁸ Effectively, because the capital stock is the sum of net investment over time, it can be argued that capital is implicitly included in the equation.

the labour force, such as children and retirees. In a reduced form specification many effects could be at work simultaneously and while clearly redundant in most production functions, the behaviour of non-working age groups affects the economy and this should be accounted for. Nonetheless, the ability of equation (1) to approximate different specifications is an empirical matter and remains to be investigated below. Regardless of whether a reduced form or structural interpretation is made, a common requirement for the appropriateness of equation (1) is that age structure and deterministic terms are sufficient to explain the trends in GDP.

3 Econometric issues

When estimating models with macroeconomic time series data in levels, such as in this paper, the properties of the series must be given careful attention. The reasons for this are twofold. First, the potential danger of spurious regressions using persistent macro data, as pointed out by Granger and Newbold (1974) and explained by Phillips (1986), warrants care in analyzing certain models. Second, the superconsistency of certain estimators when variables are cointegrated is a property that can be exploited in regressions. The main tools for analysing the time series properties of the data in this paper are the panel unit root test of Levin, Lin and Chu (2002) and two panel cointegration tests developed by Pedroni (1995, 1997a, 1999).

An assumption underlying each of the testing and estimation methods employed in the empirical section is cross-sectional independence of disturbances, meaning that all shocks are idiosyncratic. Because this is of some importance in both theoretical and empirical analysis, it will be discussed briefly here. For the process

$$\Delta \mathbf{z}_{i,t} = \boldsymbol{\varepsilon}_{i,t} \quad i = 1, \dots, N \text{ and } t = 1, \dots, T \quad (10)$$

where $\mathbf{z}'_{i,t} = (y_{i,t} \quad \mathbf{x}'_{i,t})$, $\mathbf{x}_{i,t}$ is a $M \times 1$ vector and the long-run covariance matrix of $\boldsymbol{\varepsilon}_{i,t}$ is

$$\mathbf{\Omega}_i = \sum_{j=-\infty}^{\infty} E(\boldsymbol{\varepsilon}_{i,j} \boldsymbol{\varepsilon}'_{i,0}), \quad (11)$$

cross-sectional independence requires that $E(\boldsymbol{\varepsilon}_{i,t} \boldsymbol{\varepsilon}'_{j,s}) = \mathbf{0} \quad \forall t, s \quad i \neq j$. The long-run covariance matrix for $\boldsymbol{\varepsilon}' = (\boldsymbol{\varepsilon}'_{1,t} \quad \boldsymbol{\varepsilon}'_{2,t} \quad \cdots \quad \boldsymbol{\varepsilon}'_{N,t})$ is then given by $\boldsymbol{\Sigma} = \text{diag}(\mathbf{\Omega}_1 \quad \cdots \quad \mathbf{\Omega}_N)$.⁹

While cross-sectional independence seems like an heroic assumption in most macroeconomic applications – for instance, due to correlation of business cycles across countries – it can be dealt with. The remedy most commonly suggested is to eliminate time specific effects by subtracting the cross-sectional average at time t as shown in equation (12) for the $N \times 1$ vector process \mathbf{y}_t .

$$\tilde{y}_{j,t} = y_{j,t} - N^{-1} \sum_{i=1}^N y_{i,t} \quad (12)$$

3.1 Panel unit root tests

Several different tests for unit roots in panels have been developed over the last decade, for instance by Abuaf and Jorion (1990), Maddala and Wu (1999), Hadri (2000), Levin, Lin and Chu (2002) and Im, Pesaran and Shin (2003). These panel unit root tests were to a large extent a response to the well documented low power of univariate unit root tests such as the Augmented Dickey-Fuller test (Said and Dickey, 1984).¹⁰ Among the panel unit root tests are several different specifications of null and alternative hypotheses and ways of addressing problems such as heteroscedasticity and serial correlation. Each test has its relative merits and drawbacks and the choice in this paper is on the Levin, Lin and Chu (LL) test.¹¹ Not only is its simplicity when it comes to

⁹ The definition can be extended to processes that include individual specific constants and time trends as shown by Pedroni (1995), Phillips and Moon (1999, 2000), Kao and Chiang (2000) and Levin, Lin and Chu (2002).

¹⁰ See for instance Froot and Rogoff (1995).

¹¹ For a discussion of some of the problems regarding panel unit root tests, see for instance Banerjee (1999).

estimation an appealing feature, but it has been used frequently in empirical work and the properties have been well investigated in a number of Monte Carlo studies.¹²

The LL test takes its starting point in the $N \times 1$ vector process \mathbf{q}_t and its purpose is to be able to make statements about the persistence of the included series. This is accomplished by pooling information from Augmented Dickey-Fuller (ADF) test regressions such as that in equation (13).

$$\Delta q_{i,t} = \mu_i + \eta_i t + \delta_i q_{i,t-1} + \sum_{j=1}^{f_i} \gamma_{ij} \Delta q_{i,t-j} + u_{i,t} \quad (13)$$

The specification in equation (13) is the most general and the LL test can be performed with both μ_i and η_i set to zero. Under the null hypothesis, all series have a unit root and this is tested against the alternative hypothesis of panel stationarity; formally we test $H_0 : \delta_i = 0 \quad \forall i$ versus $H_1 : \delta_1 = \dots = \delta_N < 0$. The alternative hypothesis is restrictive in the sense that it requires all series to converge towards equilibrium at the same speed.¹³ In comparison, the Im, Pesaran and Shin (2003) test allows different degrees of persistence under the alternative hypothesis, but if we believe that an economic variable has the same property regardless of country, the LL specification is preferred. When testing the null hypothesis, the test offers several test statistics and in this paper we consider the most commonly used, namely the t_{δ^*} . The test statistic is asymptotically standard normal under the null hypothesis and the null hypothesis is rejected in the left tail of the distribution.

The LL test assumes cross-sectional independence of the error terms and this assumption is, as mentioned above, unlikely to be fulfilled. Cross-sectional dependence in the panels is therefore a potential problem and O'Connell (1998) and Maddala and Wu (1999) report that the LL test has size distortions whenever cross-correlations of error terms are not controlled for. This fact was confirmed by Bornhorst (2003), who on

¹² For its use in empirical work, see Bernard and Jones (1996), Song and Wu (1997) and Bohl (1999). For Monte Carlo details, see O'Connell (1998), Maddala and Wu (1999) and Bornhorst (2003).

¹³ For a detailed description of the procedure, see Levin, Lin and Chu (2002) or Banerjee (1999).

the other hand also showed that cross-sectional dependence need not be a serious problem for these two tests if the correlations are of fairly equal size. Applying cross-sectional demeaning to the process \mathbf{q}_t , as in equation (12) above, the LL test can then be applied to the adjusted series according to

$$\Delta \tilde{q}_{i,t} = \mu_i + \eta_i t + \delta_i \tilde{q}_{i,t-1} + \sum_{j=1}^{f_i} \gamma_{ij} \Delta \tilde{q}_{i,t-j} + u_{i,t}. \quad (14)$$

Cross-sectional demeaning of the series could affect the data in such a way that the specification of deterministic in the test should be reconsidered. If, for instance, $\eta_i = \eta \forall i$, the trends in the variables could be completely removed by the demeaning and the test should be run with no time trend included. Running regressions with different assumptions regarding deterministic should therefore be a relevant robustness check.

3.2 *Panel cointegration tests*

When there is evidence of unit roots in the data, a cointegration approach can be useful in order to safeguard against spurious regressions and to detect long-run relationships. Just as was the case for the panel unit root tests, panel cointegration tests have the appealing feature of being able to increase power compared to single equation cointegration tests; for example, the Augmented Engle-Granger test (Engle and Granger, 1987) has been shown to have low power to alternatives local to unity and pooling data could therefore be useful.¹⁴ A number of tests have been suggested using quite different frameworks, such as Pedroni's (1995, 1997a, 1999) residual based tests of no cointegration, McKoskey and Kao's (1998) residual based test of the null of cointegration and Larsson and Lyhagen's (1999) likelihood based test for panel vector autoregressions. This paper will use two tests proposed by Pedroni – the group ADF and

¹⁴ For studies of the behaviour of the Augmented Engle-Granger test in a single equation framework, see Kremers et al (1992) and Haug (1996).

panel ADF tests¹⁵ – that have been frequently used in empirical work.¹⁶ While Larsson and Lyhagen’s framework is appealing – being a panel generalisation of Johansen’s (1988) methodology – it is not feasible here for to have sufficient degrees of freedom for estimation requires small N and very large T .

The group ADF and panel ADF tests take their starting point in the system of equations shown in (15) below, in which both fixed effects and individual cointegrating vectors are permitted. This flexibility can partly explain the popularity of the test, as can the fact that endogeneity/reverse causality is allowed for; exogeneity of the regressors is an assumption that is likely to be violated in many applications, including the one in this paper.¹⁷ An implicit assumption in Pedroni’s work is that there is only one cointegrating vector in each cross-section and a certain normalisation that makes sense. In this application, the issue of normalisation is straightforward since GDP is the obvious left hand side variable. Another underlying assumption that Pedroni’s tests share with the panel unit root tests, is that of cross-sectional independence of error terms. Since the solution to the problem used in this paper is the same as for the panel unit root test – that is, cross-sectional demeaning of the variables – the test is applied to cross-sectionally demeaned variables. Consequently, it is then sensible to run the regression with different specifications for the trend component in case the trend is removed by the demeaning. Specifically, we consider

$$\tilde{y}_{i,t} = \delta_i + \phi_i t + \gamma_{1i} \tilde{x}_{1i,t} + \dots + \gamma_{Mi} \tilde{x}_{Mi,t} + \varepsilon_{i,t} \quad i = 1, \dots, N \text{ and } t = 1, \dots, T \quad (15)$$

where M is the number of regressors excluding deterministic terms. The potential cointegrating relationships are individually estimated and the residuals from these

¹⁵ The small sample performance regarding size and power are investigated in Pedroni (1997a). Both tests are found to have high power – the group ADF test the highest – in a Monte Carlo study. The size distortions tended to be larger for the group ADF test than the panel ADF test in general. Other tests were found to have smaller size distortions than the group ADF and panel ADF tests in some cases but the loss in power – especially in panels with a relatively small T , such as the application in this paper – makes the group ADF and panel ADF tests seem like the preferred options.

¹⁶ See Neusser and Kugler (1998), Butler and Dueker (1999), Canning and Pedroni (1999), Edmond (2001), Sarantis and Stewart (2001) and Bahmani-Oskooee et al (2002).

¹⁷ The issue of potential endogeneity of regressors should not be overstated either on the other hand since population, rather than labour, is used in the regressions. This should reduce the endogeneity problem compared to “classic” production function estimation.

regressions are then tested for unit roots where the parameter of interest is ρ_i in equation (16).

$$\hat{\varepsilon}_{i,t} = \rho_i \hat{\varepsilon}_{i,t-1} + \sum_{j=1}^{j_i} \pi_{i,j} \Delta \hat{\varepsilon}_{i,t-j} + u_{i,t} \quad (16)$$

The group ADF and panel ADF tests differ slightly in their formulation of alternative hypothesis. For the panel ADF test, the null and alternative hypotheses are $H_0 : \rho_i = 1 \forall i$ and $H_1 : \rho_1 = \dots = \rho_N < 1$. The group ADF test on the other hand tests $H_0 : \rho_i = 1 \forall i$ versus $H_1 : \rho_i < 1 \forall i$. After the relevant transformations and adjustments, both tests have test statistics that follow standard normal distributions.¹⁸ Under the alternative hypothesis of cointegration, the test statistics diverge to negative infinity and the null is therefore rejected for observed values far in the left tail of the distribution.

3.3 *Panel estimation*

When dealing with variables that are integrated of order one, but not cointegrated, regressions such as equation (1) can be spurious in a single equation framework. In the panel case, however, spurious regression is not necessarily a problem. The intuition for this is that the cross section adds information which leads to a stronger overall signal than is the case for a single equation time series regression. For the cross-sectionally independent system defined in (10) and (11), where Ω_i is decomposed into

$$\Omega_i = \begin{bmatrix} \Omega_{y_i y_i} & \Omega_{y_i x_i} \\ \Omega_{x_i y_i} & \Omega_{x_i x_i} \end{bmatrix} \text{ and where } E(\Omega_i) = \begin{bmatrix} \Omega_{yy} & \Omega_{yx} \\ \Omega_{xy} & \Omega_{xx} \end{bmatrix},$$

Phillips and Moon (1999) showed that the pooled OLS estimator is consistent even when regressing independent random walks on each other in a panel framework. Consider estimation of the equations in (17) using data generated according to the assumptions of the above defined system.

¹⁸ For a detailed description of the procedures see Pedroni (1995, 1999).

$$y_{i,t} = \gamma_1 x_{1i,t} + \dots + \gamma_M x_{Mi,t} + \varepsilon_{i,t} \quad i = 1, \dots, N \text{ and } t = 1, \dots, T \quad (17)$$

This can equivalently be written as

$$\mathbf{y} = \mathbf{X}\boldsymbol{\gamma} + \boldsymbol{\varepsilon}$$

where \mathbf{y} is the $NT \times 1$ vector of stacked observations. The pooled OLS estimator of $\boldsymbol{\gamma}$ is then given by

$$\hat{\boldsymbol{\gamma}} = (\mathbf{X}'\mathbf{X})^{-1} \mathbf{X}'\mathbf{y} \quad (18)$$

and Phillips and Moon (1999) showed that the estimator of $\boldsymbol{\gamma}$ in (18) is consistent for the average long-run regression coefficients $\boldsymbol{\Omega}_{xx}^{-1}\boldsymbol{\Omega}_{xy}$. It should be noted that in the case of heterogeneity of cointegrating vectors, these coefficients are different from the average cointegrating coefficients.¹⁹ For the case in which the system has been generated by a more elaborate data generating process, Phillips and Moon (1999, 2000) showed that consistent estimation of parameters can also be achieved in the presence of both fixed effects and deterministic trends.²⁰ This is, however, not straight forward to show and this paper will therefore only draw on these results and not present them.

As the panel estimation also requires cross-sectional independence, relying on cross-sectionally demeaned data is more reliable. In case the deterministic trends in equation

¹⁹ Since in general $E(\boldsymbol{\Omega}_{x_i y_i}^{-1} \boldsymbol{\Omega}_{x_i y_i}) \neq [E(\boldsymbol{\Omega}_{x_i x_i})]^{-1} E(\boldsymbol{\Omega}_{x_i y_i})$, we would not expect the limits of these parameters to coincide.

²⁰ Whilst shown to be consistent estimators, the pooled OLS and fixed effects estimators are still biased as shown by for instance Phillips and Moon (1999, 2000). This bias – even though likely to be minor – raises the question of other estimators. An obvious alternative is to simply correct for the bias, but the bias corrected estimator does not improve over the standard OLS estimator in general as shown by Chen et al (1999). Other available estimators include Fully Modified OLS (FMOLS) and Dynamic OLS (DOLS) (see for instance Kao and Chiang, 2000), both of which are consistent. It is, however, shown in a Monte Carlo study by Kao and Chiang (2000) that FMOLS estimation does not improve the small sample properties over OLS in general. According to the same Monte Carlo study, there appears to be some gains in the estimation of the cointegrating vector from using DOLS as the bias is smaller. The behaviour of the DOLS estimator, however, differs depending on the number of leads and lags used in estimation. How to choose leads and lags is not obvious and further research is needed in order to understand this according to Kao and Chiang (2000). Taken together it is concluded that for the purpose in this paper the standard OLS estimator is a reasonable choice and we have the advantage that the properties both in terms of benefits and drawbacks are well known.

(1) are restricted to be homogeneous – or are very similar in the different countries – then cross-sectional demeaning will not only remove business cycle effects, but also trends. However, if there is considerable heterogeneity in the deterministic trends this might not be completely appropriate; Phillips and Moon (2000) then propose the use of OLS detrending using a linear trend and intercept and this method will also be used. Since we still require the error terms to be cross-sectionally independent, the detrended data are then cross-sectionally demeaned.

It can finally be pointed out that even if we do not risk spurious regression in the panel framework, it is still of interest to establish whether there is cointegration between the variables in the regression or not. Cointegration implies that there is a meaningful long-run relationship between the variables in the model; if the variables on the other hand are completely unrelated, we only expect the parameter vector to converge to zero.

4 Empirical study

In order to investigate the relationship between age structure and GDP, a panel consisting of 20 OECD countries with yearly observations from 1970 to 1999 is used, yielding a panel with $N = 20$ and $T = 30$. The countries in the sample are presented in Table 1. Germany is a notable exclusion due to the 1990 reunification and associated measurement problems implied by this.

Table 1. OECD countries included in the panel.

Australia	Austria	Belgium	Canada	Denmark
Finland	France	Holland	Ireland	Italy
Japan	New Zealand	Norway	Portugal	Spain
Sweden	Switzerland	Turkey	U.K.	U.S.

Real GDP data are taken from the OECD database and expressed in billions of U.S. dollars at current prices and current PPPs. Current PPPs have the advantage that spatial comparisons within a given year are straightforward as volumes are measured with the same price structure. Over time, however, comparisons incorporate effects from both relative volume changes, changes in relative prices between countries and changes in

definitions and methodologies. The drawbacks, however, are judged smaller than the advantages of the next best alternative, constant PPPs.²¹ Data on the age structure of each country are taken from the UN “*World Population Prospects: The 2000 Revision*” and given as population in thousands by age group.

4.1 *Time series properties and cointegration tests*

Turning to the time series properties of the data, all variables in the model are tested for unit roots using the LL test. The log level of GDP is a variable that is known to be trending and we are, hence, interested in testing the null hypothesis of a unit root versus trend stationarity. Regarding the number of people in each age group, it is less clear what underlying deterministic to expect; it is common in macroeconomics to assume that population is increasing over time according to some trend and some age groups – especially the retirees – appear to be doing this. However, other age groups appear to have either no trend or a downward trend. Given the lack of theoretical guidance regarding the age variables, the unit root tests will therefore be performed both with and without a trend for the log levels.

Since the tests require the series to be cross-sectionally independent, they are therefore carried out on cross-sectionally demeaned data as in equation (14). As discussed above, demeaning is another reason to justify considering different deterministic as trends in the variables may be affected by demeaning. Finally, the issue of lag length in the test equations needs to be established and again, economic theory offers little guidance. Lag length for each time series – f_i in equation (14) – is determined by performing univariate ADF tests on the individual time series for which lag length is chosen according to the Akaike (1974) information criterion. Results for the LL test applied to data in levels are presented in the top panel of Table 2. Depending on specification, the results differ slightly; the null hypothesis is not rejected in a single case when both a trend and a constant are included in the regressions. In contrast, for the specification

²¹ Like other fixed price indices, constant PPPs assumes that price structures do not change over time. It is, however, clear that relative prices do change over time and ignoring these shifts over longer periods can generate a biased picture of economic developments. Another consequence of fixing price structures at a base year is the dependence of results on the choice of the base year.

using only a constant, the null is rejected for two of the age groups. The results do not, however, offer strong evidence in favour of stationarity of the variables under consideration. To investigate whether the time series are integrated of order one – that is, stationary after first differencing – the test is next applied to the series in first differences. Results are presented in the lower panel of Table 2, where the choice of deterministic is whether to include a mean or no deterministic at all.

Table 2. Panel unit root tests.

	GDP	0-14	15-29	30-49	50-64	65+
Level						
$\mu_i \neq 0, \eta_i \neq 0$	3.613	32.386	19.698	27.834	23.875	30.515
$\mu_i \neq 0, \eta_i = 0$	3.099	6.047	-7.371**	13.946	4.085	-18.476**
First difference						
$\mu_i \neq 0, \eta_i = 0$	-11.288**	-1.318	2.955	6.407	2.915	2.183
$\mu_i = 0, \eta_i = 0$	-14.619**	-7.802**	-7.086**	-5.477**	-5.768**	-3.758**

Entries in table are the t_{δ^} test statistic from the LL test.*

Lag length in test equations is set to eight.

*** significant at the 1% level*

** significant at the 5% level*

The results shown in the lower panel of Table 2 are not straightforward to interpret. When a constant is included in the test regression, the null hypothesis is rejected only for GDP, but when the constant is omitted the test strongly supports stationarity of all the series. In general we would expect a variable that is trending in levels to have an intercept in differences, but since the data are cross-sectionally demeaned before the tests are applied to them this is not necessarily the case. In fact, it turns out that the first differenced and cross-sectionally demeaned series all have means that are very close to zero. For example, for children only three countries have series in which the mean of the first differenced series is larger than 0.01 in absolute value. The constant term in the regressions in first differences is therefore likely to be redundant, which could reduce the power of the test as the deterministic terms affect the moments used to calculate the test statistic. Hence, the tests without a constant term included are judged more reliable and the overall conclusion from the panel unit root tests is that the variables under consideration are difference stationary.²²

²² Difference stationarity of the included variables is also appealing from a theoretical point of view since it is very hard to motivate any of the variables being integrated of order two.

Concluding that the included variables appear to be generated by unit root processes, we next turn to the issue of whether they are cointegrated or not, applying Pedroni's group ADF and panel ADF tests to the data.²³ Since demeaning the data may remove any common deterministic trend in the data, equation (19) was estimated both with and without a time trend for data that have only been cross-sectionally demeaned. When data have first been detrended using a linear trend and intercept and then cross-sectionally demeaned, only the specification without a time trend was estimated. The estimation performed follows

$$\tilde{y}_{i,t} = \delta_i + \phi_i t + \gamma_{1i} \tilde{l}_{0-14i,t} + \gamma_{2i} \tilde{l}_{15-29i,t} + \gamma_{3i} \tilde{l}_{30-49i,t} + \gamma_{4i} \tilde{l}_{50-64i,t} + \gamma_{5i} \tilde{l}_{65-i,t} + \varepsilon_{i,t}$$

$$i = 1, \dots, N \text{ and } t = 1, \dots, T \quad (19)$$

where we let $\tilde{y}_{i,t}$ and $\tilde{l}_{aa-bbi,t}$ represent variables that have only been cross-sectionally demeaned as well as variables that have been detrended prior to this. The results using all three specifications are presented in Table 3.

Table 3. Panel cointegration tests.

		Cross-sectionally demeaned data	Detrended and cross-sectionally demeaned data
$\phi_i \neq 0$	Panel ADF	-5.520**	
	Group ADF	-7.233**	
$\phi_i = 0$	Panel ADF	-4.205**	-8.034**
	Group ADF	-6.594**	-11.395**

Entries in table are test statistics.

*** significant at the 1% level*

** significant at the 5% level*

Judging from the test statistics, the null hypothesis of no panel cointegration is rejected by both tests under all three specifications. That detection of cointegration supports the empirical specification and implies that a long run relationship between age structure and output exists and that a meaningful relationship can be estimated.

²³ Lag length in the individual ADF regressions – j_i in equation (16) – was determined using a step-down procedure with the maximum lag length set to four; the last lag was dropped if the absolute t -value of the corresponding coefficient was less than 1.65.

4.2 Estimation and sensitivity analysis

A fixed effects model is applied to the cross-sectionally demeaned data and a pooled OLS model to the detrended and cross-sectionally demeaned data. The models are shown in equations (20) and (21).

$$\tilde{\mathbf{y}} = \mathbf{D}\boldsymbol{\delta} + \tilde{\mathbf{L}}\boldsymbol{\gamma} + \boldsymbol{\varepsilon} \quad (20)$$

$$\tilde{\mathbf{y}} = \tilde{\mathbf{L}}\boldsymbol{\gamma} + \boldsymbol{\varepsilon} \quad (21)$$

where $\tilde{\mathbf{y}}$ is a 600x1 vector, $\mathbf{D} = [\mathbf{d}_1 \ \mathbf{d}_2 \ \dots \ \mathbf{d}_N]$ a 600x20 matrix where \mathbf{d}_i is a dummy variable indicating the i th panel member and $\boldsymbol{\gamma}$ is a 5x1 vector of slope coefficients for the previously defined age groups. The results of both estimations are presented in Table 4.

Table 4. Estimation results from panel regressions of equations (20) and (21).

	Cross-sectionally demeaned data	Detrended and cross-sectionally demeaned data
$\hat{\gamma}_1$	-0.14 (-1.22)	-0.15* (-2.12)
$\hat{\gamma}_2$	0.36** (3.61)	0.16 (1.77)
$\hat{\gamma}_3$	0.26* (1.98)	0.53** (3.78)
$\hat{\gamma}_4$	0.20* (2.11)	0.37** (3.28)
$\hat{\gamma}_5$	0.01 (0.09)	-0.08 (-0.72)
R^2	0.946	0.127
# obs	600	600

t-values in parentheses(); standard errors are Newey-West corrected.

*** significant at the 1% level*

** significant at the 5% level*

The first point to note is that the coefficient point estimates are generally in line with our expectations given the life cycle hypothesis and human capital theory: children and retirees have a negative or less positive effect on GDP than the productive age groups. The age patterns are slightly different between the two methods of detrending, but seem

fairly robust. The most surprising result is probably that in the estimation using only cross-sectional demeaning, the age group 15 to 29 has the highest point estimate. In contrast, the data that was first detrended and then cross-sectionally demeaned yielded the highest point estimate for the age group 30 to 49, more in line with the predictions of theory. One issue to be potentially concerned about is that the OLS estimator exhibits finite sample bias but given that the bias has been shown to be minor in panels for the dimensions considered here, the estimates of the parameter vector should still be fairly reliable.²⁴

Judging by the t -statistics, parameter estimates are significantly different from zero for three of the five age groups in both specifications. In both cases is the effect of the 30-49 and 50-64 year olds upon the level of GDP significant whilst the parameter estimate for 65 year olds and over insignificant. The t -statistics should be interpreted with some caution, for only when regressors are strictly exogenous and error processes homogeneous over all members of the panel are t -statistics correct. When these conditions are not met, t -statistics will only serve as approximations when testing hypotheses and this should be taken into account when making inference.²⁵ For the present application though, there is good reason to believe that inference based on the ordinary t -statistics should be reasonably reliable. Monte Carlo evidence provided by Coakley et al (2001) demonstrates that when cointegration is present, the distribution of t -statistics is close to standard normal when the parameter vector is the same over all countries, or there only is moderate heterogeneity. The high value of the R^2 seen in Table 4 also points to the validity of the regressions, as Kao (1999) has shown that the R^2 in panel spurious regressions is very low in general.

A word of caution regarding the assumption of cross-sectional independence underlying the panel estimations; the estimated contemporaneous covariance matrices of the residuals from equations (20) and (21) show no evidence of being diagonal. A number of tests – not reported, but available upon request – are performed, all of which reject

²⁴ See Kao and Chiang (2000) and Coakley et al (2001).

²⁵ See Phillips and Moon (1999) for a theoretical discussion of this issue and Kao and Chiang (2000) and Coakley et al (2001) for Monte Carlo evidence.

the null hypothesis of the two matrices being diagonal.²⁶ The effects of violating the assumption of cross-sectional independence are not completely clear. Phillips and Moon (1999) suggest that this may affect some of the asymptotic theory, but are rather unspecific and due to the only recent development of the panel cointegration methods, studies concerning empirical evidence regarding this issue are also few. The only available study, performed for panel DOLS by Mark and Sul (2003), indicates that imperfectly controlling for cross-sectional dependence does not seem to affect estimation and testing very much. Though panel DOLS is a different method to the pooled OLS and fixed effects estimators used in this study, the major modification incorporated into the panel DOLS estimator concerns endogeneity of regressors and serial correlation, not cross-sectional dependence, so there is little reason to believe that the effect should be very different here.

Based on the results from the above mentioned Monte Carlo studies regarding panel cointegration, it appears as if there should only be minor potential pitfalls regarding t -statistics and significance of regressors. However, to validate the results, the distributions of the slope coefficients were bootstrapped. The bootstrap was performed for the pooled OLS estimation – which due to its theoretically slightly more reasonable parameter estimates was judged the preferred specification – under two different assumptions. In the first bootstrap no notice was taken of the fact that the pooled OLS estimation is based on a number of countries and the randomisation was over both i and t without restrictions. NT observations were drawn with replacement and the parameters of the models then estimated; this was repeated 2000 times. The second bootstrap acknowledged the presence of individual countries and therefore randomisation took place over the N countries and if a country was chosen, all T observations for that country were included in the sample. The observations were then replaced and N countries were chosen this way. Like the first bootstrap, this procedure was repeated 2000 times. Rejection percentiles for the parameters were generated accordingly and the results from the bootstraps are shown in columns two and three of Table 5.

²⁶ The assumption of cross-sectional independence is also violated in the panel unit root tests and panel cointegration tests. For the LL panel unit root test, the effects of this are – as previously discussed – fairly well established. The effects on the panel cointegration tests are not that well examined, but Pedroni (1997b) found that empirical cross-sectional dependency does not appear particularly important in panel based tests on real exchange rates.

Table 5. Results from sensitivity analysis of equation (21).

	Bootstrap	Bootstrap – country effects	Country 1 to 10	Country 11 to 20	1970 to 1984	1985 to 1999
$\hat{\gamma}_1$	-0.15**	-0.15	-0.20* (-1.97)	-0.12 (-1.46)	-0.01 (-0.14)	-0.40** (-4.18)
$\hat{\gamma}_2$	0.16*	0.16	0.11 (1.06)	0.23 (1.62)	0.13 (1.33)	0.15 (0.88)
$\hat{\gamma}_3$	0.53**	0.53*	0.58** (3.29)	0.45* (2.36)	0.28* (2.14)	0.81** (3.49)
$\hat{\gamma}_4$	0.37**	0.37*	0.34** (2.89)	0.42** (2.68)	0.06 (0.46)	0.63** (5.30)
$\hat{\gamma}_5$	-0.08	-0.08	-0.24 (-1.39)	0.08 (0.59)	-0.21 (-1.59)	-0.01 (-0.08)
R^2	0.127	0.127	0.127	0.147	0.055	0.273
# obs	600	600	300	300	300	300

Bootstrapped parameter distributions based on 2000 replications are used in the first two columns to generate rejection percentiles – seed is set to 1208.

t-values in parentheses(); standard errors are Newey-West corrected.

** significant at the 1% level

* significant at the 5% level

Comparing the bootstrap distributions of parameters, four out of five parameters are significant at the five percent level for the “fully randomised” bootstrap, whereas for the “country effects” bootstrap only two parameters are found significant. These results can be compared to the pooled OLS estimation with robust standard errors in which three of the age groups were found to have significant explanatory power. Regarding the different bootstrap methods, it could be the case that the method using full randomisation over i and t underestimates the standard errors of the parameters as country specific heteroscedasticity and autocorrelation are not completely accounted for. The second method might, on the other hand, overestimate the standard errors. A problem with this method in small N panels – such as the one in this study – is that as the same country may appear several times in the same replication, there can be considerably less variance in dependent and explanatory variables for the bootstrap samples compared with the original sample. The overall impression is, however, that using bootstrap methods does not dramatically change the conclusions compared to the standard inference: age structure seems to have significant explanatory power on GDP and in theoretically predicted directions.

As a final sensitivity analysis of the results, we investigate whether any specific country or period drive the results. Sensitivity analysis was performed by dividing the sample in

two different ways; first, along the cross-sectional dimension, yielding two samples each consisting of ten countries over 30 years; second, along the time dimension, which generates two samples with 20 countries observed over 15 years. The results from these four sub-sample estimations are presented in columns four to seven of Table 5 above. Though parameter estimates – as well as the significance of these – and fit of the model vary over the different samples, the general pattern is robust. The point estimate is highest for the 30 to 49 year olds in all four samples and point estimates for children and the retired imply a negative or relatively small positive effect on GDP from these age groups. The robustness of these sub-sample results further strengthens the relevance of the age structure effects established in this study.

Regarding the interpretation of the coefficients, it makes most sense to see them as favouring the life cycle hypothesis and human capital theory in a general way, but they could also – as stated earlier – be given the more structural interpretation as exponents from a Cobb-Douglas aggregate production function. Under this assumption, a number of hypotheses – such as constant returns to scale and whether labour is homogeneous – could be tested. Whilst t - and F -tests are only approximations in the estimated models, the bootstrapping exercise above showed that the approximation might be reasonable in equation (21) so the hypotheses $\gamma_1 + \gamma_2 + \gamma_3 + \gamma_4 + \gamma_5 = 1$ (constant returns to scale) and $\gamma_2 = \gamma_3 = \gamma_4$ (homogeneous labour) are therefore tested. Based on F -tests – keeping in mind that the tests should be taken with a pinch of salt – the null hypothesis of constant returns to scale cannot be rejected at the five percent level whereas that null hypothesis of homogeneous labour can.²⁷ Though an aggregate production function is a possible underlying assumption for the specifications estimated above, this interpretation should not be over emphasised. An F -test of the effects of children and retirees on GDP can be seen as one way of distinguishing between a strict production function interpretation that would dictate relevance only of working age groups and a reduced form specification that includes all age groups. Performing this test – that is, testing $\gamma_1 = \gamma_5 = 0$ – it is found to reject the null hypothesis at the five percent level which

²⁷ The observed F -value for the former is 1.13 and for the latter, 15.71.

suggests that viewing the estimation results as those of a reduced form specification is probably more appropriate.²⁸

5 Conclusions

The life cycle hypothesis and human capital theory are well known concepts to all economists. Whilst they have had large impact on empirical research at the microeconomic level for many years, their influence on macroeconomics has been smaller. During the last two decades, research focusing on the relationship between macroeconomic variables and age structure has intensified and this paper aims to further this literature by investigating the relationship between GDP and population, building upon the theoretical predictions of the life cycle hypothesis and human capital theory.

Using a panel of 20 OECD countries between 1970 and 1999, there is strong support for a cointegrating relationship between GDP and the number of people in five different age groups. The estimated relationship largely supports the life cycle hypothesis and human capital theory; children and retirees are found to have negative or relatively less positive effects on GDP than productive age groups. The details of the underlying mechanisms behind these age structure effects are, however, not completely transparent and in order to improve our understanding, further research in the field is needed. It can still be concluded that the twin assumptions used in many macroeconomic models of stable population growth and the representative agent are violated by data. By acknowledging population heterogeneity and tracing its effects onto the level of GDP, the results of this paper add sophistication to standard macroeconomic modelling. They will also provide useful information to both policymakers facing particularly large or small cohorts feeding through their countries' age distribution and forecasters wishing to make medium- or long-term forecasts.

²⁸ Observed F -value is 8.96.

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