Is the Feldstein-Horioka puzzle still with us? National saving-investment dynamics and international capital mobility: A panel data analysis across EU member countries

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Abstract

This paper attempts to investigate the degree of financial integration and international capital mobility by analysing the dynamics of national saving-investment relationships. We interpret the relationship between national saving and investment in the long run as reflecting a solvency constraint and focus on the short-term saving-investment relationship to assess the degree of capital mobility. We apply the Panel Autoregressive Distributed Lag (PARDL) model proposed by Pesaran *et al.* (1999) using data for 14 EU member countries from 1970 to 2007. Our empirical results suggest that there exist a close relationship between saving and investment in the long run that is consistent with the existence of a solvency constraint that is binding for each country in the long run. We also find that the parameter for the error-correction term is always highly significant, which supports the choice of an error-correction formulation. Moreover, we show that the parameter estimated for the error-correction term, i.e. the speed of adjustment to the long-run equilibrium, varies with the sample period considered. The estimated speed of adjustment becomes smaller in absolute terms as more recent data are included in the sample, which indicates that deviations from long-run equilibrium current accounts have become more persistent over time, signalling some degree of capital mobility.

Key Words: *Capital Mobility; Feldstein-Horioka puzzle; Saving-Investment Relationship; Current account adjustment; Dynamic panel cointegration; European Union*

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1. Introduction

As a result of technological breakthroughs, financial liberalization, and growth in the volume of trade (Obstfeld, 1995), during the last three decades the degree of international financial integration, defined as the extent to which financial markets are connected, has increased substantially in both developed and developing countries, giving capital and other assets the opportunity to move unobstructed between countries. Increasing capital mobility across countries is an important phenomenon for economic policy makers and firms. It has potential beneficial effects on the economy because it enables agents to allocate the resources more efficiently and give more scope for risk management, has an influence on the effects of economic policy and responses to external shocks and it allows for investment and hence growth beyond the premises of domestic saving. As such the level of capital mobility among countries is clearly an important question in international economics. For analysts, the assumption of a high or low level of capital mobility has profound implications for their modeling strategy; for policy makers, the degree of capital mobility may significantly affect the impact of different policy instruments. In the literature of open-economy macroeconomics, defining and measuring capital mobility has been one of the most important issues.

There are two main methods of measuring the level of international capital mobility (Obstfeld, 1993). The first one involves comparing the movement of rates of return on capital across countries, a common approach when interest is in analyzing financial capital flows, while the second approach focuses on the actual international capital flows. The present paper adopts the latter approach, and in particular it examines the implications of the existence of correlation between saving and investment rates across countries for the level of capital mobility. The focus on capital flows rather than rates of return reflects an interest in whether real (as opposed to financial) capital has been mobile among economies; by contrast, studies of the behavior of relative rates of return have tended to concentrate on the behavior of financial capital. Since the focus is on long-run real capital flows, this paper dwells on the second approach. However, it should be noted that even with external financial reform the possibilities for capital flow might be limited by obstacles, such as transaction costs, taxes, and official restrictions.

In their seminar paper Feldstein and Horioka (1980) investigate the correlation of saving and investment across countries and they consider the non-existence of such a relationship as evidence of high degree of capital mobility. Their model is based on the standard goods market equilibrium equation and measures the extent to which a higher domestic saving rate is associated with a higher rate of domestic investment. If capital is indeed very mobile, the relationship between saving and investment should be weak and conversely, if capital is rather immobile, investment rates should correspond closely to saving rates. Many economists have studied the savings and investment relation [see Frankel (1992) and Coakley et al.(1998)] since the seminal work of Feldstein and Horioka.

Some of these studies focused on the time series aspects of the data. Savings and investment rates usually turn out to be non-stationary. It is well known that one should avoid a spurious regression problem by checking the co-integration relationship when the time series data are non-stationary. However, the traditional co-integration technique has the problem of low power. In order to improve the power of the test, the number of observations (span of data) should be extended. However, it is not easy to find data for a very long time span except for a few countries (exception is Taylor [1996] who uses the data of saving and investment rate for 12 countries over the period 1850-1992). Furthermore, expansion of the time horizon might cause the unwanted regime-shift problem for the saving-investment relationship. The panel data enable us to solve the power problem. Panel data, besides providing more information, gives greater power and less size distortions than the standard time series unit root and co-integration tests [see Levin *et al*. (2002)]. Furthermore, the ability of panel data estimation techniques to acknowledge cross-section specific effects would also yield better and more representative estimation results. These fixed or random effects may represent unobservable factors like different economic policies being followed, different capital control measures and any other time invariant country-specific factor that is not easily observable but may still be significant in determining the saving-investment relationship.

In light of the level and volatility of capital inflow experienced by many European countries during the period 1970-2000, an investigation of the validity of the FH-puzzle is highly warranted.

Therefore, the present study addresses the FH issue in the case of the EU14 countries over the period 1970-2007. In many respects, this group of countries presents an interesting sample for empirical investigation. For instance, to the extent that country size influences the degree of capital mobility, the EU14 includes both small and large economies with different levels of development. Furthermore, they exhibit different economic structures, different degrees of integration in the international economy, and different growth performances over the years, and thus different profit opportunities for international capital. In the empirical tests that follow, allowance is also made for the effects of different exchange rate regimes. Furthermore, the present paper is distinctive in that it employs recently developed panel co-integration techniques to account for cross-sectional dependence. We also aim at checking if FH coefficients using panel co-integration approach is a good candidate for measuring the international capital mobility. A few papers have examined international capital mobility using panel co-integration techniques with data only on developed countries [see Ho (2002) and Moon and Phillips (1998)]. To the best of our knowledge, there is no research on FH coefficient in European countries, using panel co-integration techniques.

This paper is distinct from other researches in several respects. First, it applies the recently developed panel co-integration techniques to the relationship between savings and investment in Asian countries. Furthermore this paper uses the 'between-group' (or group-mean) panel co-integration test to the fully modified OLS (hereafter FMOLS) and dynamic OLS (hereafter DOLS) methods in order to deal with heterogeneity problems and to conduct plausible tests. Second, this study measures the international capital mobility of the European economy as a group, rather than for individual countries.

The paper is organized as follows. Section 2 provides a theoretical background to the FH hypothesis and reviews the methodologies used to test it as well as alternative interpretations and critiques. Section 3 presents the model and the econometric methodology, while Section 4 describes the data and discusses the empirical results from panel data analyses whereas the summary and concluding remarks are given in the final section.

2. Literature review

2.1 The Feldstein-Horioka approach: theory critique and alternative interpretations

Feldstein and Horioka (1980) proposed a simple model based on the goods market equilibrium condition in an attempt to explain the degree of capital mobility. The paper measures the extent to which a higher domestic savings rate is associated with a higher rate of domestic investment. With perfect capital mobility, the relationship between saving and investment should be very weak. Conversely, if capital is rather immobile, investment rates should correspond closely to saving rates. In other words they argued that increased financial integration should decrease the correlation between domestic investment and saving rate. The investment rate for country i can be written as follows:

$$\left(\frac{I}{Y}\right)_i = \gamma_i - \delta r_i + \varepsilon_i \qquad (1)$$

where I is a measure of gross domestic investment, Y is the gross domestic product, r the domestic real interest rate, γ the intercept, and ε represents all other factors that determine investment. Since it is assumed that the national saving rate is a function of the real interest rate, Feldstein and Horioka estimate the following equation:

$$\left(\frac{I}{Y}\right)_i = \alpha_i + \beta\left(\frac{S}{Y}\right)_i + \nu_i \qquad (2)$$

where S is gross domestic saving measured as gross domestic product minus private and government consumption, α the intercept, and v the error term. The series for investment and saving are scaled by GDP as a simple way of controlling for business cycle fluctuations on estimates of β. Following Feldstein and Bacchetta (1990), β, can be referred to as the "saving retention coefficient"-later called the FH coefficient- as it reveals the extent to which an increase in domestic savings finances domestic investment. They [Feldstein and Horioka (1980), Feldstein (1983) and Feldstein and Bacchetta (1989)] interpret the positive correlation coefficient as an evidence of a low degree of the long-term international capital mobility. A high correlation coefficient means that investments have been financed mainly by domestic savings. On the contrary, if capital mobility is high the correlation coefficient should be low since investments might be financed by savings from abroad. The reason behind this is that capital moves from the countries where it is less efficient to those where it is more efficient (Hogendorn, 1998, p.142). As the degree of mobility increases, higher portion of domestic savings would be invested elsewhere in the world. When national savings and investment rates are equal to each other, the current account balance will be close to zero. This means that investments made by domestic residents are matched by their own savings. It is important to note that provision is made for open economies in the national goods market equation, and that saving and investment need not be equal in a specific period within a country because of international capital flows. This implies that it would be highly unlikely that the coefficient β would be exactly equal to unity, *i.e.* there is no movement of capital between countries. In order for β to equal zero, three conditions are necessary: real interest parity must hold, the world real interest rate must be exogenous or in no way correlated with the saving rate, and there must be no correlation between the saving rate and ε.

When estimating the equation for a sample of 16 industrial countries for the period 1960-74, Feldstein and Horioka found the estimated β coefficient to be in the range 0.85 to 0.95 and to be insignificantly different from unity indicating that just 5-15% of national savings was invested abroad. These estimated coefficient values contradicted the prior expectations of near perfect capital mobility in the selected OECD countries, especially because of the fact that this period was characterized by the many efforts made by countries to enhance the interaction of global capital markets. Although perfect capital mobility may be perceived in the short run, there appeared to be sufficient elements, rigidities and preferences to keep saving invested in the country of origin.

2.2. The intertemporal budget constraint approach: The Long-run relationship

Using the average cross-sectional data across 16 OECD countries for the period from 1960 to 1974 Feldstein and Horioka (1980) on their controversial paper find the estimated “saving-retention coefficient” ranging from 0.87 to 0.90. They concluded that some 90% of domestic savings remains within a country to finance domestic investment. Capital, therefore, is not internationally mobile, in contradiction of the belief that the industrialized countries have few barriers to capital movements.

Drawing on the existing literature on public sector’s solvability analysis (Hamilton and Flavin, 1986; Trehan and Walsh, 1988. 1991; Kremers, 1989), it is argued in this paper that a regression of investment on saving should not be interpreted as a measure of capital mobility, but as empirically capturing the effects of the intertemporal budget constraint of a country, which in the long-term is an indicator of its solvability. These solvency constraint arguments have also been used to explain the relationship between savings and investment as both a short-run and a long-run phenomenon (Finn, 1990; Sinn, 1992; Baxter and Crucini, 1993; Coakley *et al.*, 1996; Coakley and Kulasi, 1997; Jansen, 2000). In this way, domestic saving and investment should be perfectly correlated in the long-run, since the current balances i.e. the differences between savings and investment, should add up to zero. In an open economy, due to the intertemporal budget constraint, current account deficits/surpluses cannot be sustained indefinitely. Therefore, in the long-run investment cannot deviate too much from savings. In this way, when a country borrows or lends in the short term and when it makes the opposite operations at a subsequent date, respect of the intertemporal budget constraint induces the existence of a long-term relation between national saving and domestic investment. This implies that saving and investment should keep a one-to-one relation in the steady-state, in spite of their temporary divergence in the short-term from their long-term equilibrium levels. In other words, savings and investment are cointegrated variables and this implies a state of current account constancy. In the long-term current account constancy induces a strong correlation between domestic saving and investment and this is interpreted in the Feldstein and Horioka approach as evidence of imperfect capital mobility. On the other hand, in the short-run, the size and sign of the correlation between savings and investment would depend on the structure of the economy, as well as on the nature of the shocks. So, while small positive, zero or negative correlations would suggest a significant degree of capital mobility, high positive correlations would not necessarily imply imperfect capital mobility.

Coakley *et al*. (1996), however, argue that a no Ponzi game condition or what Buiter and Patel (1992) have called weak solvency in the context of government deficits, generates a Feldstein and Horioka unity coefficient. Therefore, the high saving-investment correlation in the long-run must be interpreted more as reflecting the intertemporal budget constraint of a country than a measure of international capital mobility (Husted, 1992; Jansen, 1996; Jansen and Schulze, 1996; Moreno, 1997).

When the rest of the world agrees to hold an increasing share of a country's external debt, the country can maintain a greater real absorption level than its domestic real income. Thus, a current account deficit can be regarded as a consequence of the net capital inflows corresponding to lower saving than domestic investment. The more capital is mobile internationally, the more saving transfers are internationally responsible for the external imbalances. Current account deficits, however, can only be durably financed if the country is considered as solvable in the long-run. Public sector's solvability can then be extended to the notion of the external finance of a country (Hakkio and Rash, 1991; Husted, 1992; Coakley *et al*., 1996, 1997, 1998, 1999, 2004; Moreno, 1997). Following Hakkio and Rush (1991) the one-period budget constraint of a country can be written as:

$$I_t + (1 + r_t)B_{t-1} = S_t + B_t \tag{3}$$

where B_t represents the funds raised by issuing new debt or external borrowing, r_t is the (one-period) interest rate, I_t is domestic investment, and S_t corresponds to national saving. Rearranging terms eq. (3) can be written as:

$$I_t + B_{t-1} + r_t B_{t-1} = S_t + B_t$$

or

$$I_t - S_t + r_t B_{t-1} = B_t + B_{t-1}$$

and

$$I_t - S_t + r_t B_{t-1} = \Delta B_t \tag{4}$$

Eq. (4) implies that the difference between saving and investment and the accumulated interest payments to foreigners in the previous period are financed by increases in external borrowing. Using a forward solution in eq. (4) as proposed by Hakkio and Rush (1991) yields:

$$B_0 = \sum_{t-1}^{\infty} \mu_t (S_t - I_t) + \lim_{n \to \infty} \mu_n B_n \tag{5}$$

where

$$\mu_t = \prod_{s=1}^{t} \beta_s, \beta_s = 1/(1 + r_s) \text{ and } \mu_n = \prod_{s=0}^{n} \beta_s$$

The crucial element in the intertemporal budget constraint is the last term $\lim(\mu_n B_n)$ where the limit is taken as $n \to \infty$. As discussed by Hamilton and Flavin (1986), Barro (1987), and McCalum (1984), the limiting value of $(\mu_n B_n)$ must equal zero in order to rule out the possibility of financing the repayment of the debt by issuing new debt, that is excludes the possibility of using this sort of Ponzi scheme to finance its debt (Moreno, 1997) In order to derive empirically testable implications we need an alternative equation by transforming eq. (5). Assuming that the world interest rate is stationary, with unconditional mean equal to *r*, eq. (5) can be expressed as:

$$Z_t + (1 + r)B_{t-1} = S_t + B_t \tag{6}$$

where $Z_t = I_t + (r_t - r)B_{t-1}$

Solving eq. (6) forward as proposed by Hakkio and Rush(1991) gives:

$$B_{t-1} = \sum_{j=0}^{\infty} \beta^{j+1}(S_{t+j} - Z_{t+j}) + \lim_{j\to\infty} \beta^{j+1} B_{t+j} \qquad (7)$$

where $\beta = 1/(1+r)$ and Δ is the first difference operator. Eq. (7) can be rewritten as:

$$I_t + r_t B_{t-1} = S_t + \sum_{j=0}^{\infty} \beta^{j-1}(\Delta S_{t+j} - \Delta Z_{t+j}) + \lim_{j\to\infty} \beta^{j+1} B_{j+1} \qquad (8)$$

The left hand side of (8) represents investment expenditure as well as interest payments due to the country's foreign debt. If we subtract S_t from both sides of (7) and multiply by minus one each side, then the left hand side of (8) can be expressed as the economy's current account (Husted, 1992).

Assuming that S and Z are both non-stationary processes, and each is I(1) i.e. integrated of order one, then S_t and Z_t can be expressed as random walks with drift:

$$S_t = \alpha_1 + S_{t-1} + u_{1t} \qquad (9)$$

$$Z_t = \alpha_2 + Z_{t-1} + u_{2t} \qquad (10)$$

where α_j (j=1,2) are drift parameters (possibly equal to zero) and the terms u_{jt} are stationary processes. In this case (8) can be rewritten as:

$$S_t = \alpha + R_t - \lim_{j\to\infty} \beta^{j+1} B_{j+1} + u_t \qquad (11)$$

where $R_t = I_t + r_t B_{t-1}$, $\alpha = \sum \beta^{j-1}(\alpha_2 - \alpha_1) = [(1+r)/r][\alpha_2 - \alpha_1]$ and $u_t = \sum \beta^{j-1}(u_{2t} - u_{1t})$.

Equation (11) forms the basis of the hypothesis tests in this paper. Assuming that the limit term in (11) equals zero when $j \to \infty$, then (11) can be transformed into a standard regression equation as:

$$S_t = \alpha + bR_t + e_t \qquad (12)$$

where under the null hypothesis that the country is satisfying its intertemporal budget constraint, we would expect that $b = 1$ and e_t would be stationary. Thus, is S and R are non-stationary processes, then under the null hypothesis of solvability there will exist a linear stationary combination of S and R, and they are thus cointegrated. In this case, cointegration is a necessary condition for the country to be obeying its intertemporal budget constraint. As Hakkio and Rush (1991) demonstrate the condition that $b=1$ does not. In the case where B_{t-1}, the initial external debt, is positive, then b need only be less than or equal to one for the constraint to hold. However, if, say, S is non-stationary while R is stationary, there is no long-run relationship between S and R. This implies that the country is violating its intertemporal budget constraint because S tends to grow while R does not. Therefore, the limit term, for which we derive an expression (eq. (11)) does not equal to zero. This violates the intertemporal budget constraint. The current account, represented by the difference between national saving and investment, is stationary in the long term. Under these conditions, the existence of a saving and investment relationship in the long term reflects the respect of the country's intertemporal budget constraint and involves one-to-one adjustment between these two variables. From an empirical point of view, the long-run relationship between saving and investment rates as implied by the intertemporal budget constraint amounts to testing not just such a long-run relationship exists, but also the unit coefficient cointegration hypothesis (Blanchard and Fischer,1989) that accords with the standard sready-state (Miller, 1988; Argimon and Roldan, 1994).

3. Econometric approach

3.1. Panel unit root tests

Before proceeding to cointegration techniques, we need to verify that all variables are integrated with the same order. In doing so, we have used the first generation panel unit root tests of Levin *et al*. (2002), Im *et al*. (2003), Breitung (2000), the Fisher-type tests using ADF and PP tests of Maddala and Wu (1999) and Hadri (2000) as well as the second generation panel unit root tests of Pesaran (2005) and Moon and Perron (2004). From the relevant tables (1 and 4) we conclude that the null hypothesis of the unit roots for the panel data of the investment and saving series cannot be rejected in level. Therefore, we can apply tests for panel cointegration between savings and investment.

3.2. Panel cointegration tests

Once the stochastic properties of the variables have been determined, we apply Kao's (1999) and

Pedroni's (1995 and 1999) panel cointegration test methodology. Kao (1999) tests the residuals $\hat{e}_{it}$ of the OLS panel estimation by applying DF- and ADF - type cointegration tests as follows:

$$\hat{e}_{it} = \hat{e}_{it}\rho_{it-1} + v_{it} \tag{1}$$

and

$$\hat{e}_{it} = \hat{e}_{it}\rho_{it-1} + \sum_{j=1}^{p}\phi_j \Delta\hat{e}_{it-j} + v_{it} \tag{2}$$

The null hypothesis of no co-integration, H_0 is $\rho = 1$, is tested against the alternative hypothesis of stationary residuals, H_1: $\rho < 1$. The OLS estimate of ρ can be written as:

$$\hat{\rho} = \frac{\sum_{i=1}^{N}\sum_{t=2}^{T}\hat{e}_{it}\hat{e}_{it-1}}{\sum_{i=1}^{N}\sum_{t=2}^{T}\hat{e}_{it}^2} \tag{3}$$

Kao (1999) has developed five DF and ADF types of cointegration tests using panel data, the asymptotic distributions of which converge to a standard normal distribution N(0, 1). The test statistics are DF-roh* , DF-t* and ADF, which are for cointegration with the endogenous regressors, and DF-roh and DF-t which are based on assuming strict endogeneity of the regressors.

Pedroni (1995) suggests a Phillips – Perron type panel co-integration test, which implies less strict assumptions with respect to the distribution of the error terms than do the DF and ADF tests described above. Pedroni (1995) provides two test statistics, PC1 and PC2, which converge to a standard normal distribution. First, under the null hypothesis of no co-integration, the panel autoregressive coefficient estimator $\hat{\rho}_{N.T}$ can be constructed as follows:

$$\hat{\rho}_{N,T-1} = \frac{\sum_{i=1}^{N}\sum_{t=2}^{T}\hat{e}_{it-1}\Delta\hat{e}_{it} - \hat{\lambda}_i}{\sum_{i=1}^{N}\sum_{t=2}^{T}\hat{e}_{it-1}^2} \tag{4}$$

where $\widehat{\lambda}_i$ acts as a scalar equivalent to the correlation matrix, Γ, and corrects for any correlation effect. Pedroni (1999) provides the limiting distribution of two test statistics:

$$PC_1 = \frac{T\sqrt{N(\widehat{\rho}_{N,T-1})}}{\sqrt{2}} \Rightarrow N(0,1) \tag{5}$$

$$PC_2 = \frac{\sqrt{NT(T-1)}}{\sqrt{2}}(\widehat{\rho}_{N,T-1}) \Rightarrow N(0,1) \tag{6}$$

Table 5 shows the results of the cointegration tests between savings and investment rates. The results indicate that the null hypothesis of no cointegration between saving and investment ratios can be rejected at conventional significance levels in all cases.

Similarly to the Im *et al*. (2003) and Maddala and Wu (1999) panel unit root, the panel cointegration tests proposed by Pedroni (1999) also take into account heterogeneity by employing parameters that are allowed to vary across individual members of the sample. Taking into account such heterogeneity constitutes an advantage because it is unrealistic to assume that the vectors of cointegration are identical among individuals on the panel.

The first step of the implementation of Pedroni's (1999) cointegration test requires the estimation of the following long-run relationship:

$$y_{it} = \alpha_i + \delta_i t + \beta_i x_{it} + \varepsilon_{it} \tag{7}$$

for i=1,…., N, t=1,….T, where N refers to the number of individual members in the panel and T refers to the number of observations over time. The structure of estimated residuals is as follows:

$$\widehat{e}_{it} = \widehat{\rho}_i \widehat{e}_{it-1} + \widehat{u}_{it} \tag{8}$$

Pedroni (1999) proposed seven different statistics to test panel data cointegration. Out of these seven statistics, four test statistics are based on pooling, what is referred to as the ‘‘Within’’ dimension, and the last three are based on the ‘‘Between’’ dimension. No co-integration is the maintained hypothesis for both these group of tests statistics. However, the distinction among them is derived from the specification of the alternative hypothesis. For the tests based on ‘‘Within’’, the alternative hypothesis is $\rho_i = \rho < 1$ for all ***i***, while with respect to the last three test statistics, which are based on the ‘‘Between’’ dimension, the alternative hypothesis is $\rho_i < 1$ for all i.

The finite sample distribution for these seven statistics has been tabulated by Pedroni (1999) with the use of Monte Carlo simulations. The calculated statistics must have values that are smaller than the tabulated critical value to reject the null hypothesis of the absence of cointegration.

3.3. Panel cointegration estimation

We use a panel error-correction approach to analyze the long-run relationship separately from the short-run adjustment, estimating long- and short-run effects jointly from a general autoregressive distributed-lag (ARDL) model suggested by Pesaran *et al*. (1996, 1999). The approach is an extension of the time-series analysis and pooling approach used by Jansen (1996, 1998). It drops the homogeneity assumption implied by the pooling approach in the latter, combining both cross- section and time-series analyses to more fully exploit the cross-country variation in the data. Specifically, we use three different techniques. In addition to the dynamic fixed-effects (DFE) estimator, we use the pooled mean group estimator (PMG) and the mean group estimator (MGE). The methods differ in the extent to which they allow for cross-country heterogeneity of parameter estimates. At one extreme, the fully heterogeneous coefficient model, MGE, imposes no cross-country constraints and is estimated on a country-by-country basis. Coefficient estimates are obtained as the unweighted mean of the estimated coefficient for the individual countries. At the other extreme is the fully homogeneous coefficient model, the DFE estimator, which imposes both the long-run and short-run coefficients and error variances to be the same between cross- section units of the panel. It allows only the intercepts to differ across countries. The PMG estimator can be interpreted as an intermediate procedure between the DFE and MGE, since it involves a mixture of pooling and averaging: it assumes that the long-run relationship is similar across countries, but it allows the short-run coefficients to vary across countries.

The choice among these estimators is a trade-off between consistency and efficiency. In effect, estimators that impose restrictions dominate the heterogeneous models in terms of efficiency if the restrictions are valid. In particular, if the long-run coefficients are equal across countries, then the PMG will be consistent and efficient, whereas the MGE will only be consistent. If the long-run restrictions are wrongly imposed, however, the DFE and PMG estimates will be inconsistent, and estimates of the speed of adjustment will be biased downwards (Robertson and Symons, 1992; Pesaran and Smith, 1995).

It is difficult to say a priori which method among the three approaches is more appropriate in the context of the sample under examination. OECD countries form a relatively homogeneous group of countries, especially in terms of the extent to which they have liberalized their capital accounts. This would suggest that the long-run relationship between saving and investment rates is similar across these countries. Nonetheless, some impediments remain to capital mobility, and countries differ substantially in terms of size and economic structure, so that, even within the group of relatively homogeneous EU member countries, differences in the long-run relationship between saving and investment rates could persist. Against the background of these considerations, we apply all three different approaches, and evaluate their performance based on our assessment of the plausibility of results and statistical tests of homogeneity of error variances and of short- and long-run slope coefficients.

The unrestricted specification for the system of the PARDL equations for t=1,2,…T and i=1,2,…,N is :

$$y_{it} = \sum_{j=1}^{p} \theta_{ij} y_{i,t-1} + \sum_{j=1}^{q} \gamma_{ij}^{'} x_{i,t-j} + \mu_i + \varepsilon_{it} \qquad (9)$$

where $x_{i,t-j}$ is the (k x1) vector of explanatory variables for group *i* and μ_i are the fixed effects. In principle the panel can be unbalanced and *p* and *q* may vary across countries. Equation (9) can be re-parameterised as a VECM system.

$$\Delta y_{it} = \alpha_i + \lambda_i (y_{i,t-1} - \beta_i^{'} x_{i,t-1}) + \sum_{j=1}^{p-1} \delta_{ij} \Delta y_{ij,t-1} + \sum_{j=1}^{q-1} \delta_{ij}^{'} \Delta x_{i,t-j} + \mu_i + \varepsilon_{it} \qquad (10)$$

where the β_i are the long-run parameters and λ_i are the equilibrium (or error) correction parameters.

The DFE, MGE and the PMG estimators for the whole panel are given by:

$$\beta'_{DFE} = \beta'_i, \lambda_{DFE} = \lambda_i, \delta'_{DFE} = \delta'_i, \forall_i$$

$$\beta'_{MGE} = (1/N)\sum_{i=1}^{N} \beta'_i, \lambda_{MGE} = (1/N)\sum_{i=1}^{N} \lambda_i, \delta'_{MGE} = (1/N)\sum_{i=1}^{N} \delta'_i$$

$$\beta'_{PMG} = \beta'_i, \forall_i, \lambda_{PMG} = (1/N)\sum_{i=1}^{N} \lambda_i, \delta'_{PMG} = (1/N)\sum_{i=1}^{N} \delta'_i$$

Estimation could be conducted by applying iterated least squares, imposing and testing the cross-country restrictions on β. However, this will be inefficient as it ignores the contemporaneous residual covariances. A natural estimator is Zellner's SUR method (Zelner, 1962), which is a feasible GLS approach. SUR estimation is only possible if N is smaller than T. Thus PSS suggest a maximum likelihood estimator. In our case SUR is feasible, but we use the efficient Pesaran *et al.* (1999) method. The MG estimation derives the long run parameters for the panel from an average of the long run parameter from PARDL models for individual countries (see Pesaran and Smith, 1995). The PMG method of estimation, introduced by Pesaran *et al.* (1999) occupies an intermediate position between the MG method, in which both slopes and the intercepts are allowed to differ across country, and the standard fixed effects method, in which the slopes are fixed and the intercepts are allowed to vary. In PMG estimation, only the long run coefficients are constrained to be the same across countries, while the short run coefficients are allowed to vary.

Given equation (2), the steady-state equilibrium (in country i) can be defined as follows:

$$\alpha_i + \lambda_{DFE}(y_i - \beta'_{DFE} x_i) = 0$$

$$\alpha_i + \lambda_{PMG}(y_i - \beta'_{PMG} x_i) = 0$$

$$\alpha_i + \lambda_{MGE}(y_i - \beta_{MGE}^{'} x_i) = 0, \forall_i = 1,...., N, t = 1,...., T$$

Provided that a long-term saving investment is significantly different from (1, −1), the error correction coefficient, $\hat{\lambda}$, can be interpreted as an indicator of capital mobility, with a lower $\hat{\lambda}$ indicating higher capital mobility. On the other hand, if the cointegrating vector is not statistically different from (1, −1) no conclusion can be drawn.

Pesaran *et al*. (1999) argue that in panels omitted group specific factors or measurement errors are likely to severely bias the country estimates. It is a commonplace in empirical panel studies to report a failure of the 'poolability' tests based on the group parameter restrictions. So Pesaran *et al*. (1999) propose a Hausman test. This is based on the result that an estimate of the mean long-run parameters in the model can be derived from the average (mean group) of the country regressions. This is consistent even under heterogeneity. However, if the parameters are in fact homogeneous, the mean and the individual parameters coincide and the PMG estimates are more efficient. Thus we can form the test statistic

$$H = \hat{q}^{'}[\mathrm{var}(\hat{q})]^{-1}\hat{q} \sim x_k^2 \qquad (11)$$

where $\hat{q}$ is a (k x 1) vector of the difference between the mean group and PMG estimates and $\mathrm{var}(\hat{q})$ is the corresponding covariance matrix. Under the null that the two estimators are consistent but one is efficient, *var(q)* is easily calculated as the difference between the covariance matrices for the two underlying parameter vectors. If the poolability assumption is invalid then the PMG estimates are no longer consistent and the test rejects.

3.4. Breitung's (2005) two-step estimator

[see tables 12,13 and 14 Alternative estimation results for long-run coefficient]

Furthermore, cross-sectional dependence appears to be a recent but vital issue in dynamic panel estimation. When there is correlation across units, it is necessary to consider the correlation in the estimation. The literature to remedy this problem is still growing. One of the attempts to handle cross sectional dependence is Breitung (2005).

His model is the panel analogue of Johansen methodology. He proposes a two-step estimator for the estimation of long run cointegrating vector.

To explain this procedure, consider a VECM model

$$\Delta y_{it} = \alpha_i \beta' y_{i,t-1} + \varepsilon_{it}, i = 1, 2, ...N; t = 1, 2, ...T \qquad (1)$$

where εit is a k-dimensional white noise error vector with $E(\varepsilon_{it}) = 0$ and positive definite covariance matrix $\sum_i = E(\varepsilon_{it}.\varepsilon_{it-1})$. The long-run parameter β can be obtained conditional on some consistent initial estimator of α_i and $\sum_i$. To derive the two-step estimator, Breitung (2005) transforms (1) into a VECM below

$$\gamma_i' \Delta y_{it} = \gamma_i' \alpha_i \beta' y_{i,t-1} + \gamma_i' \varepsilon_{it} \qquad (2)$$

$$z_{it} = \beta' y_{i,t-1} + v_{it} \qquad (3)$$

where $z_{it} = (\gamma_i' \alpha_i)^{-1} \gamma_i' \Delta y_{it}, v_{it} = (\gamma_i' \alpha_i)^{-1} \gamma_i' \varepsilon_{it}$ and γ is a k x r matrix with $rk(\gamma_i' \alpha_i) = r$ .It follows that $\sum_v -(\alpha_i' \sum^{-1} \alpha_i)^{-1}$ is positive semi-definite and , therefore, the optimal choice of the transformation is $\gamma_i' = \alpha_i' \sum_i^{-1}$ .The resulting estimator is asymptotically equivalent to the Gaussian ML estimator.

In the first stage, the individual equation of (2) is estimated by VAR and conducts a nor- malization procedure to obtain α_i and $\sum_i$. The restriction that the cointegrating vectors are the same for all cross-sectional units is ignored, but this does not affect the asymptotic properties of the estimator. For the asymptotic properties of the two-step estimator, it is only required that the parameters are estimated consistently as T approaches infinity.

At the second estimation stage, the system is transformed into (3) such that the cointe- gration matrix β can be estimated by least-squares of the pooled regression.

4. Empirical results

4.1 Unit Root tests

We perform several panel unit root tests for the variables in question. Results of panel unit root tests differ as we allow for more heterogeneity and variations in the tests. For all tests, expect for that of Hadri the null hypothesis is the non-stationarity of the series. So, high *p*-values imply unit root in the series. First, we start with the first generation panel unit root tests, and then the second generation panel unit root tests are utilized.

In Table 1 we report the results from the application of first generation panel unit root tests, which fail to allow for general forms of cross-sectional dependence. Remarkably, neither the panel *t*-statistics of Levin, Lin and Chu (2002) and Im, Pesaran and Shin (1997, 2003) nor the inverse chi square of Maddala and Wu (1999) and Choi (2001) are able to reject the joint non-stationarity null, thus favouring the unit root hypothesis, while we are able to reject the null of panel stationarity in the case of Hadri panel stationarity test. More specifically,

The first test Levin, Lin and Chu (2002) assume the homogeneity of the coefficient of the lagged dependent variable in the alternative hypothesis and no cross-sectional dependence. Results of this ADF-type test show that the null hypothesis of unit root is not rejected for both the variables considered. Results are robust regardless of the choice of bandwidth parameters (Table 1).

The second test Im, Pesaran and Shin (1997, 2003) is again based on cross- sectional independence assumption but allows for heterogeneity in coefficients of lagged dependent variables in the alternative hypothesis. Instead of pooling the data like Levin, Lin and Chu (2002), they test each unit separately for nonstationarity.. According to the test, saving and investment can be considered as nonstationary (Table 1).

The third test is discussed in two different papers with similar approach. Maddala and Wu (1999) and Choi (2001), both of them derive statistics from individual p-values of unit root tests for each cross-sectional unit. They test the same hypothesis as of Im, Pesaran and Shin (1997, 2003). Results are in favor of unit root for both variables at 1 percent significance level. (Table 1).

After the first-generation unit root test, the second-generation unit root tests, in which cross-sectional dependence assumption is relaxed, are applied to our dataset. To examine dependence across units, we use a simple tests proposed by Pesaran (2004).Table 2 contains CD statistics that employ residuals from ADF estimations with intercept only. The hypothesis of zero cross section correlation is rejected and ADF specifications at the 1% level of significance. The outcomes of these tests clearly indicate the presence of cross-sectional dependence for both variables in question. Consequently, tests for the presence of unit roots in the our panel data should take this dependence into account to produce unbiased and reliable results. Tables 3,4, 5 and 6 present the results from the application of the second generation panel unit root tests of Chang (2002), Breitung and Das (2005), Moon and Perron (2004), Bai and Ng (2004), Choi (2006), Phillips and Sul (2003), Harris *et al*. (2005) and Pesaran (2007).

To examine dependence across units, the most influential method in the literature is the factor structure approach. The results for the Phillips and Sul(2003,PS) panel test are reported in Table 3. The inverse normal Z statistic strongly implies that the unit root hypothesis should not be rejected. This meta statistic combines the *p*-values of the individual ADF regressions of the de-factored data. The highly significant Z test statistic is a reflection of the failure to reject the unit root hypothesis for nearly all individual series, which have been removed from the common factor in the first step. To remedy dependence problem Pesaran(2007) proposes cross-sectionally augmented Dickey Fuller test statistic (CADF) by adding cross-section average of lagged levels and first-differences of the individual series to conventional Dickey Fuller or augmented Dickey Fuller regressions. He also proposes a truncated version of CADF (CADF*) to avoid extreme outcomes that may arise in small time dimensions. His test statistics are cross-sectional average of CADF and CADF*. In fact, they are just the cross-sectionally augmented version of the statistics offered by Im, Pesaran and Shin (1997, 2003) denoted as CIPS and CIPS*, respectively. Results, reported in Tables 3 and 6, should be interpreted according to optimal lag length. The CIPS and CIPS* statistics are not smaller than any of the critical values corresponding to the 1%, 5% and 10% level of significance. Otherwise, the outcomes are not very sensitive to the choice of number of lagged differences. Thus, on the basis of the common unobserved factor assumption for the error process, the Pesaran test gives indication of non-stationarity of our panel data. The results of the panel stationarity test of Harris et.al. (2005) that allows for at least one factor as reported in Table 3 strongly reject the null stationarity in our panel data with the

$\hat{S}_k^F$ statistic. So far, the tests of Pesaran (2007, Phillips and Sul (2003) and Harris et.al.(2005) failed to reject the unit root hypothesis for our panel data when allowing a single factor structure in the composite error term. Whether the one-factor model suffices to assess the dependence structure of the panel data will ascertained by the following approaches. Table 3 also reports the outcomes of the procedure from the Moon and Perron (2004,MP) test. As discussed in previous sections, the application of the Moon and Perron approach requires the estimation of the number of common factors. In implementing this test, the seven information criteria for estimating the number of factors that are proposed by Bai and Ng (2002) were considered. In the application of the information criteria to the saving and investment rates in the balanced EU 14 countries panel, congruent results obtained when setting $k^{\max} = 6$ and focusing on BIC_3 in which case it is recommended to assume one common factor. The latter is the preferred criterion of Moon and Perron in small samples. Notice that under the assumption of the common factor, the data generating processes of the Moon and Perron and Phillips and Sul tests are the same, and the only difference lies in the treatment of the common unobserved factor in the estimation strategy. As in the paper by Moon and Perron, the long-run variances were estimated using the Andrews and Monahan (1992) method. The t_{α}^{*} statistic implies that the null of a homogenious unit root in the panel for the assumption of one, two or six common factors should not be rejected as indicated in Table 3. In contrast, the t_b^{*} statistic rejected the null for all specifications. When considering the results of both test statistics, the conclusions to be drawn are highly contradictory. However, none of the simulations studies in Moon and Perron test (see Gengenbach et.a.,2004; Gutierrez, 2006) provides guidance as to which *t* statistic should be preferred in applications. Thus, the Moon and Perron test offers no clear conclusion in the present analysis.

The most comprehensive study in this respect is Bai and Ng (2004). Factor structure is based on the idea of decomposing a variable into two unobserved components, common factor and idiosyncratic error. The former is strongly correlated with many of the series and the latter is largely unit specific. Accordingly, for a series to be nonstationary either the idiosyncratic error or some of the common factors should be nonstationary. Bai and Ng (2004) propose testing common factors and unit specific shocks separately. For the number of common factors equal to one, the statistic they offer is a version of $ADF_{\hat{F}}^{c}$ test statistic.

For the number of common factors greater than one, their statistics, which are corrected (MQ_c) and filtered test (MQ_f), give the number of independent common stochastic trends. If the number of common independent stochastic trends is equal to zero, then there are N cointegrating vectors for N common factors, and that all common factors are stationary. For idiosyncratic errors, they propose a test statistic defined as in Choi (2001).

Results of Bai and Ng (2004) test, reported in Table 4 are in favor of unit root for our variables. Idiosyncratic shocks to each variale are all nonstationary. For investment, the number of common factors is one and its p-value is 0.395 implying nonstationarity. For saving, the number of common factors is equal to the number of common independent stochastic trend, i.e. at least two independent nonstationary common factors can be identified.

For robustness purposes, we also conduct three additional tests for nonstationarity. One of them is unit root test in Choi (2002). This test also allows for cross-sectional dependence and has a specification based on error component model. Results of this test are again in favour of unit root. Finally, the tests of Breitung and Das (2005) and Chang (2002) clearly fail to reject the joint nonstationarity hypothesis. At this point, a caveat applies. The nonlinear IV panel unit root test of Chang (2002) has been recently criticised by Im and Pesaran (2003) since under strong forms of cross-correlation, the Chang tests may display size distortions. This shortcoming may be alleviated in panels with large T relative to N as in our case. Also, even if size distortions existed, we fail to reject the joint nonstationarity null, thus supporting the view that saving-investment rates are best described as a unit root.

4.2 Cointegration tests

The cointegration analysis is carried out for our panel data using Kao (1999) and Pedroni (1995,1999) approaches and the results are presented in Tables 7 and 8. using Kao (1999) and Pedroni (1995) panel cointegration tests are in essence an application of the Engle and Granger (1987) cointegration analysis. As in the analysis of single time series, these approaches test the residuals from the estimation for stationarity. Kao (1999) and Pedroni (1995) provide different statistics for this purpose, both of which assume homogenous slope coefficients across countries. Table 7 shows the outcomes of the cointegration tests between savings and investment rates. The results indicate that the null hypothesis of no cointegration between savings and investment ratios can be rejected at conventional significance levels in all cases. Table 8 shows the outcomes of Pedroni's (1999) cointegration tests between the investment and savings rates. We use four within-group tests and three between-group tests to check whether the panel data are cointegrated. The columns labelled within-dimension contain the computed value of the statistics based on estimators that pool the autoregressive coefficient across different countries for the unit root tests on the estimated residuals. The columns labelled between- dimension report the computed value of the statistics based on estimators that average individually estimated coefficients for each country. The results of the within-group tests and the between-group tests show that the null hypothesis of no cointegration cannot be rejected at the 1% significance level. Therefore, the ratios of savings and investment are cointegrated for the panel of all countries and for the panels of country groups.

The presence of a long-run relationship between the investment and savings rates in the panel of EU countries is economically meaningful in that it suggests that these countries meet the long-run solvency condition. Having found that there exists a cointegrating link between the two variables (ratios of savings and investment), it is convenient that the savings retention coefficient be estimated using a panel cointegrating estimator. In this paper, we choose to employ the Mean Group and the Pooled Mean Group estimator (Pesaran et al.,1995;1999).

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Table 1 : Panel Unit Root Tests

Series	t^*_{LLC}	t_{IPS}	MW_λ ADF	MW_λ PP	Breitung t- Test	Hadri Z-Test
			Fisher x^2			
(I/Y)	1.82571 (0.9661)	-1.43204 (0.0761)	33.8325 (0.2064)	29.0085 (0.4121)	-1.24619 (0.1063)	6.29009* (0.0000)
(S/Y)	0.75823 (0.7758)	1.57989 (0.9429)	17.5813 (0.9298)	19.8646 (0.8694)	-.60381 (0.2730)	7.74914* (0.0000)
Δ(I/Y)	-15.2485* (0.0000)	-15.2353* (0.0000)	229.001* (0.0000)	318.079* (0.0000)	-14.9661* (0.0000)	2.22404 (0.0131)
Δ(S/Y)	-10.8044* (0.0000)	-9.32676* (0.0000)	143.764* (0.0000)	256.637* (0.0000)	-6.53394* (0.0000)	3.58399 (0.0000)

Note: p-values, obtained from Eviews 7.1, in parentheses. (*) shows rejection of the null hypothesis. All the tests are conducted with individual effects and a linear trend. Lags=4 for both levels and first-differences. For the MW_λ test, the degrees of freedom are 2N(=28).

Table 2: Results of the Pesaran's CD tests

Investment Rate		
Deterministic Terms	Lags	CD statistic
Intercept	0	18.03*
	1	19.41*
	2	18.91*
	3	18.39*
Intercept and Ttrend	0	17.86*
	1	19.17*
	2	18.59*
	3	17.69*
Savings Rate		
Deterministic Terms	Lags	CD statistic
Intercept	0	10.97*
	1	10.03*
	2	9.90*
	3	8.94*
Intercept and Ttrend	0	12.35*
	1	10.29*
	2	9.84*
	3	9.19*

* Significance at the 1% level, in rejecting the null-hypothesis

Table 3: The Pesaran CADF unit-root test results

Investment Rate		
Countries	With Intercept $CADF_i$	With Intercept and Trend $CADF_i$
Austria	-1.916	-2.009
Belgium	-1.462	-2.433
Denmark	-2.369	-3.091
Finland	-1.467	-2.295
France	-1.591	-1.184
Germany	-1.506	-3.467*
Greece	-1.675	-2.586
Ireland	-1.822	-2.296
Italy	-3.417*	-2.693
Netherlands	-2.189	-2.341
Portugal	-1.335	-1.550
Spain	-0.734	-2.322
Sweden	-1.417	-2.032
UK	-1.873	-1.978
Savings Rate		
Countries	With Intercept $CADF_i$	With Intercept and Trend $CADF_i$
Austria	-2.053	-1.649
Belgium	-1.407	-2.212
Denmark	0.069	-3.252
Finland	-2.062	-2.057
France	-4.174*	-3.913*
Germany	-1.546	-1.189
Greece	-2.638	-0.655
Ireland	-1.411	-1.176
Italy	-2.218	-3.233
Netherlands	-1.863	-1.472
Portugal	-0.365	-3.752*
Spain	-2.080	-2.272
Sweden	-1.111	-1.815
UK	-1.163	-3.020

Lags=3

For $CADF_i$, with intercept and intercept and trend the 5% critical values are -3.28 and -3.79 respectively (Pesaran,2007,Table I(b) and I(c)

Table 4: Second-generation panel unit-root tests*

Investment Rate	Pesaran's CIPS	Pesaran's CIPS*	Moon and Perron t_{-a}	t_{-b}
With Intercept	-1.769	-1.769	-0.257	-0.344
With Intercept and Trend	-2.306	-2.306	-1.088	-1.504
Savings Rate	Pesaran's CIPS	Pesaran's CIPS*	Monn and Perron t_{-a}	t_{-b}
With Intercept	-1.716	-1.716	-1.335	-1.616
With Intercept and Trend	-2.303	-2.303	-1.552	-1.267

*For levels with lags=3

For CIPS with intercept and intercept and trend the 5% critical values are -2.16 and -2.65 respectively (Pesaran,2007 Table II(b) and II(c).

Table 5 : Panel Homogeneous[a] Co-integration Tests

Kao(1999)[b]					*Pedroni*(1995)[c]	
DF-rho	DF-t	DF-rho*	DF-t*	ADF	PC_1	PC_2
-5.5820 (0.0000)	-3.4951 (0.0000)	-13.4162 (0.0000)	-3.9285 (0.0000)	-3.2769 (0.0000)	-20.2632 (0.0000)	-19.9948 (0.0000)

Note: P-values are in parenthesis.

a. Homogeneous tests assume equality of all coefficients in the null and alternative hypothesis.
b. The DF test statistics are analogous to the parametric Dickey-Fuller test for non-stationary time series. The DF-rho and DF-t statistic assume strict exogeneity of the regressors with respect to errors and no autocorrelation. DF-rho* and DF-t* statistics are based on endogenous regressors. These tests depend on consistent estimates of the long-run variance-covariance matrix to correct for nuisance parameters once the limiting distribution has been found. The ADF test is analogous to the ADF test for non-stationary time series.
c PC1 and PC2 are two non-parametric versions of the Phillips-Perron tests.

Table 6 : Pedroni's Homogeneous [a] panel cointegration tests

Test Statistic	Statistic	Probability
Within-dimension		
Variance ratio statistic	2.115878	0.0062
Rho statistic	-3.066456	0.0011
Phillips-Perron statistic	-3.660566	0.0001
ADF *t*-statistic	-5.111756	0.0000
Between-dimension		
Rho statistic	-3.899971	0.0001
Phillips-Perron statistic	-2.402310	0.0081
ADF *t*-statistic	-3.983163	0.0000

Notes

a. Homogeneous tests allow all the coefficients in the null and alternative hypothesis, to differ across countries under the alternative hypothesis.

The *within*-dimension test statistics are based on estimators that effectively pool the autoregressive coefficient across different members for the unit root tests on the estimated residuals. The *between*-dimension test statistics are based on estimators that average the individually estimated coefficients for each country. The first of the four statistics is a type of non-parametric variance ratio statistic. The second is a panel analogue of the Phillips-Perron (1988) rho-statistic. The third is a panel analogue of the Phillips-Perron (1988) *t*-statistic. The fourth is a panel analogue of the Dickey-Fuller *t*-statistic. The asymptotic distribution of each of the four statistics is normal with zero mean and unit variance. As such the standard normal table provides the critical values.

Table 7 : Panel error-correction estimates of saving-investment relationship: Common parameters, Unobserved common component, EU-14, 1970-2007

Estimators	1970-2007		
	DFE	PMG	MGE
Long-run coefficients			
Constant		1.384***	0.879
		(5.418)	(1.195)
Saving rate	0.838***	0.784***	1.06***
	(8.331)	(11.769)	(6.063)
Adjustment term			
Error-correction term	-0.221***	-0.279***	-0.294***
	(-8.021)	(-7.441)	(-7.142)
Short-run coefficients			
Δ (Saving rate)	0.336***	0.310***	0.283***
	(6.307)	(5.051)	(4.747)
Memorandum term			
Hausman's test		2.91	
p-Value		(0.09)	

Notes: The *t*-value in the parentheses. (*), (**), (***) denote the 10, 5 and 1 per cent significance level respectively.
a. The dynamic fixed-effect estimator (DFE) allows short-run coefficients. The pooled mean group estimator(PMG) allows short-run coefficients, including the adjustment term and the variances of the error term to differ across countries, while the long-run saving rate coefficient is constraint to be the same. The mean group estimator (MGE) allows the long-run saving rate coefficient to differ between countries as well. For all estimators, an ARDL(1,1) specification is used, where the first number and the second number in parenthesis stand for the lag length of the lagged dependent and the explanatory variable, respectively.
b. Hausman's test determines the validity of the assumption made for long-run saving rate coefficient across countries (i.e. comparing PMG and MGE estimation results.

Table 7a. Panel error-correction estimates of saving-investment relationship: Common parameters, Unobserved common component, EU-14, 1970-1989

Estimators	1970-1989		
	DFE	PMG	MGE
Long-run coefficients			
Constant		1.512*** (4.040)	1.118 (0.789)
Saving rate	0.852*** (8.555)	0.932*** (20.015)	0.582*** (5.437)
Adjustment term			
Error-correction term	-0.371*** (-7.390)	-0.522*** (-6.387)	-0.588*** (-6.326)
Short-run coefficients			
Δ (Saving rate)	0.368*** (4.283)	0.103 (0.818)	0.028 (0.184)
Memorandum term			
Hausman's test		0.66	
p-Value		(0.42)	

Table 7b: Panel error-correction estimates of saving-investment relationship: Common parameters, Unobserved common component, EU-14, 1990-2007

Estimators	1990-2007		
	DFE	PMG	MGE
Long-run coefficients			
Constant		1.239*** (5.188)	4.439 (1.213)
Saving rate	0.701*** (5.945)	0.744*** (7.477)	0.170* (1.812)
Adjustment term			
Error-correction term	-0.348*** (-7.696)	-0.322*** (-4.998)	-0.394*** (-4.641)
Short-run coefficients			
Δ (Saving rate)	0.046 (0.738)	0.159*** (2.292)	0.169*** (2.381)
Memorandum term			
Hausman's test		0.93	
p-Value		(0.34)	

Table 8: Individual results of PMG and MGE estimates of saving-investment relatioship,1970-2007

Countries	Pooled mean group estimates PMG^a			Mean group estimates MGE^a			
	Adjustment term	Δ(Saving Rate)	Adjusted R^2	Saving rate	Adjustment term	Δ(Saving Rate)	Adjusted R^2
Austria	-0.493*** (-3.476)	0.262 (1.193)	0.362	0.803*** (5.125)	-0.493*** (-3.285)	0.261 (1.120)	0.363
Belgium	-0.459*** (-3.195)	0.085 (0.251)	0.288	0.572*** (3.065)	-0.558*** (-3.122)	-0.061 (-0.160)	0.309
Denmark	-0.192** (-2.157)	0.479** (2.287)	0.287	0.577 (1.489)	-0.210** (-2.089)	0.486** (2.196)	0.292
Finland	-0.107 (-1.641)	0.607*** (3.158)	0.386	0.310*** (2.950)	-0.101 (-1.321)	0.510*** (2.698)	0.492
France	-0.145* (-1.703)	0.782*** (3.595)	0.418	1.306** (3.056)	0.187* (-1.958)	0.699*** (2.955)	0.441
Germany	-0.105 (-1.486)	0.407*** (2.705)	0.348	1.272** (2.558)	-0.128 (-1.651)	0.373*** (2.313)	0.367
Greece	-0.469*** (-3.054)	0.216 (1.484)	0.362	0.679*** (5.092)	-0.477*** (-2.955)	0.236 (1.533)	0.375
Ireland	-0.156** (-2.219)	0.105 (0.481)	0.195	1.443 (1.059)	-0.116 (-1.235)	0.121 (0.515)	0.107
Italy	-0.431*** (-3.259)	0.180 (0.655)	0.269	0.991*** (5.248)	-0.461*** (-3.292)	0.128 (0.441)	0.294
Netherlands	-0.358*** (-4.252)	-0.005 (-0.039)	0.351	0.638*** (3.134)	-0.361*** (-4.078)	0.012 (0.094)	0.362
Portugal	-0.267** (-2.579)	0.786* (1.800)	0.252	0.552** (2.141)	-0.326** (-2.475)	0.377* (1.673)	0.265
Spain	-0.239** (-2.176)	0.015 (0.061)	0.186	1.257*** (2.464)	-0.232* (-2.464)	-0.004 (-0.001)	0.117
Sweden	-0.183** (-2.056)	0.364** (2.362)	0.396	0.832** (2.092)	-0.176* (-1.161)	0.371** (2.159)	0.396
UK	-0.301*** (-2.623)	0.454** (2.175)	0.345	0.844** (2.535)	-0.293** (-2.325)	0.457** (2.064)	0.346

Notes: The *t*-values in the parentheses.
a The PMG constraints the coefficient of the saving rate to be the same for all countries. It is estimated to be equal to 0.789, with a *t*-value of 13.56. In the case of both PMG and MGE, an ARDL (1, 1) specification is used.

Table 9a : Robustness of Estimates to Country Size, 1970-2007

Size of *GDP*[1]	Pooled Mean Group estimates(PMG)		
	Large	Medium	Small
Long-run coefficients			
Constant	-0.132 (-0.789)	2.286*** (3.739)	1.322*** (3.238)
Saving rate	1.039*** (7.090)	0.684*** (7.816)	0.809*** (6.356)
Adjustment term			
Error-correction term	-0.254*** (-8.568)	-0.354*** (-5.483)	-0.241*** (-3.576)
Short-run coefficients			
Δ (Saving rate)	0.427*** (3.492)	0.128* (1.804)	0367*** (4.871)
Memorandum term			
Hausman's *test*[2]	0.74	1.66	1.11
p-Value	(0.39)	(0.20)	(0.29)

Notes: The *t*-value in the parentheses. * ,**, *** denote the 10,5 and 1 per cent significance level respectively. We performed separate panel regressions for each group of countries similar to Kim(2001).

1. Groups are selected according to the size of the sample -average real GDP dominated in US dollars. Accordingly, Large countries are France, Germany, Italy and UK. Medium countries are Belgium, Greece, Netherlands, Spain and Sweden. Small countries are Austria, Denmark, Finland, Ireland and Portugal.
2. Hausman's test determines the validity of the equality of the long-run saving rate coefficients across EU countries by comparing the PMG and MGE estimates.
3. Groups are selected according to the size of the non-traded sector, which is approximated, following Wong(1990), by the sample-average of the export minus the imports over GNP ratio. Large countries are Italy, Portugal and Spain. Medium countries include Finland, France, Germany, Greece, Sweden and UK. Small countries are Austria, Denmark, Belgium, Ireland and Netherlands.

Table 9b : Robustness of Estimates to Country Size, 1970-2007

Size of *GDP*[1]	Mean Group estimates(MGE)		
	Large	Medium	Small
Long-run coefficients			
Constant	-0.200 (-0.350)	2.000* ((1.849)	0.622 (0.359)
Saving rate	1.103*** (9.890)	0.795*** (6.468)	1.289*** (2.722)
Adjustment term			
Error-correction term	-0.267*** (-3.653)	-0.361*** (-5.029)	-0.249*** (-3.415)
Short-run coefficients			
Δ (Saving rate)	0.414*** (3.150)	0.111* (1.345)	0.351*** (4.843)
Memorandum term			
Hausman's *test*[2]			
p-Value			

Table 9c : Robustness of Estimates to Openness, 1970-2007

Size of non-traded sector[3]	Pooled Mean Group estimates(PMG)		
	Large	Medium	Small
Long-run coefficients			
Constant	0.690*** (3.795)	0.977*** (2.204)	2.067*** (4.546)
Saving rate	0.967*** (6.267)	0.803*** (7.452)	0.701*** (6.974)
Adjustment term			
Error-correction term	-0.309*** (-3.994)	-0.217*** (-3.854)	-0.341*** (-4.874)
Short-run coefficients			
Δ (Saving rate)	0.176* (1.416)	0.471*** (5.893)	0.177* (1.965)
Memorandum term			
Hausman's test	0.06	2.95	0.67
p-Value	(0.80)	(0.12)	(0.41)

Table 9d : Robustness of Estimates to Openness, 1970-2007

Size of non-traded sector	Mean Group estimates(MGE)		
	Large	Medium	Small
Long-run coefficients			
Constant	1.284 (0.763)	-0.312 (-0.233)	2.066*** (2.689)
Saving rate	0.933*** (4.545)	1.334*** (3.679)	0.806*** (4.806)
Adjustment term			
Error-correction term	-0.340*** (-5.113)	-0.227*** (-3.998)	-0.348*** (-4.186)
Short-run coefficients			
Δ (Saving rate)	0.167 (1.496)	0.441*** (6.882)	0.164* (1.685)
Memorandum term			
Hausman's test			
p-Value			

Table 10 : Estimation of the Saving rate coefficient with Static Panel Estimators

	1970-2007	1970-1989	1990-2007
Between[1]	0.527*** (11.09)	0.521*** (16.67)	0.234*** (4.16)
Pooled[1]	0.587*** (18.06)	0.619*** (16.62)	0.302*** (5.61)
Within[1]	0.575*** (12.08)	0.749*** (10.16)	0.353*** (6.17)
Random effects[1]	0.612*** (18.75)	0.712*** (17.24)	0.372*** (7.35)
Memorandum term			
Hausman's *test*[2]	0.173	1.636	0.320
p-Value	(0.68)	(0.200)	(0.571)

Notes: White heteroskedasticity-consistent standard errors. *t*-values in parentheses.
*** denotes the 1% level of significance.
1.The between estimator is obtained from the average value of each country and therefore emphasizes the inter-country dimension. The pooled estimator assumes individual homogeneity as well as the temporal stability of the relation. The within estimator introduces heterogeneity through individual fixed effects and is calculated from the difference between the saving and investment ratio and the individual average of the variable. The random effects model introduces heterogeneity through a specific unobservable country effect in the error term.
2. Hausman's test checks for fixed vs. random effects.

Table 11: Panel error-correction estimates of saving-investment relationship: Common parameters Unobserved Common Component, EU-14

Estimators	1970-2007		
	DFE	PMG	MGE
Long-run coefficients			
Constant		2.875***	2.116***
		(9.191)	(2.061)
Saving rate	0.716***	0.745***	0.954***
	(7.374)	(11.400)	(7.309)
Adjustment term			
Error-correction term	-0.256***	-0.387***	-0.402***
	(-8.06)	(-12.054)	(-10.094)
Short-run coefficients			
Δ (Saving rate)	0.329***	0.227***	0.188
	(6.183)	(4.031)	(2.746)
Ttrend	-0.018***	-0.030***	-0.030*
	(-2.187)	(-2.301)	(-1.920)
Memorandum term			
Hausman's test		3.42	
p-Value		(0.06)	

Table 12: Alternative estimation results for Long-run Coefficient

Panel I{1970-2007)					
	Static Fixed Effect	Dynamic Fixed Effect(DFE)	Mean Group Estimation (MGE)	Panel Mean Group Estimation (PMG	Panel 2-Step Estimation (Breitung)
Saving rate	0.599***	0.838***	1.060***	0.784***	0.445***
t-statistic	(17.478)	(8.333)	(6.063)	(11.769)	(6.313)

Note: ***, **.* denote significance at the 1, 5, and 10% level, respectively.

Table 13: Alternative estimation results for Long-run Coefficient

Panel II{1970-1989)					
	Static Fixed Effect	Dynamic Fixed Effect(DFE)	Mean Group Estimation (MGE)	Panel Mean Group Estimation (PMG	Panel 2-Step Estimation (Breitung)
Saving rate	0.727***	0.852***	0.582***	0.932***	0.642***
t-statistic	(15.842)	(8.544)	(5.457)	(20.015)	(13.843)

Note: ***, **.* denote significance at the 1, 5, and 10% level, respectively.

Table 14: Alternative estimation results for Long-run Coefficient

Panel III {1990-2007)					
	Static Fixed Effect	Dynamic Fixed Effect(DFE)	Mean Group Estimation (MGE)	Panel Mean Group Estimation (PMG	Panel 2-Step Estimation (Breitung)
Saving rate	0.345***	0.701***	0.170*	0.744***	0.375***
t-statistic	(6.780)	(5.945)	(1.812)	(7.477)	(7.361)

Note: ***, **.* denote significance at the 1, 5, and 10% level, respectively.

Figure 1: Short-run adjustment speed by country

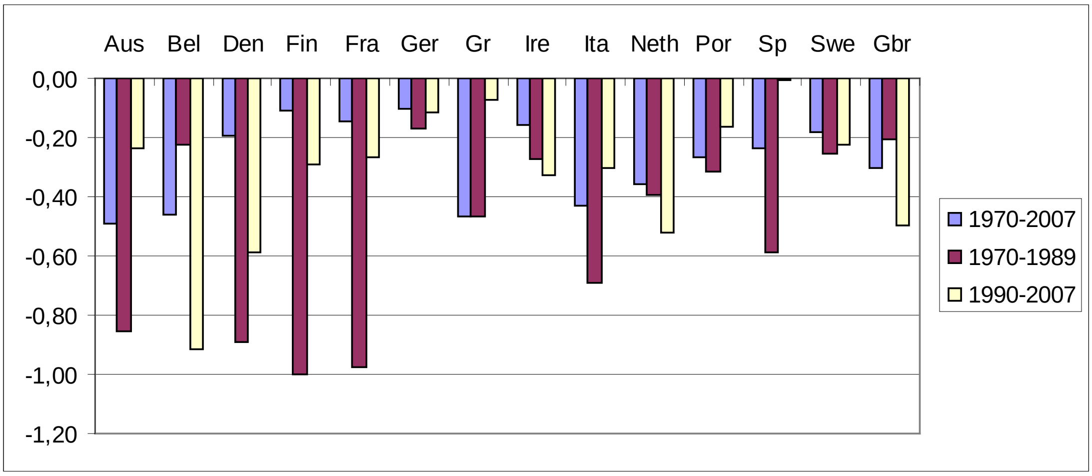
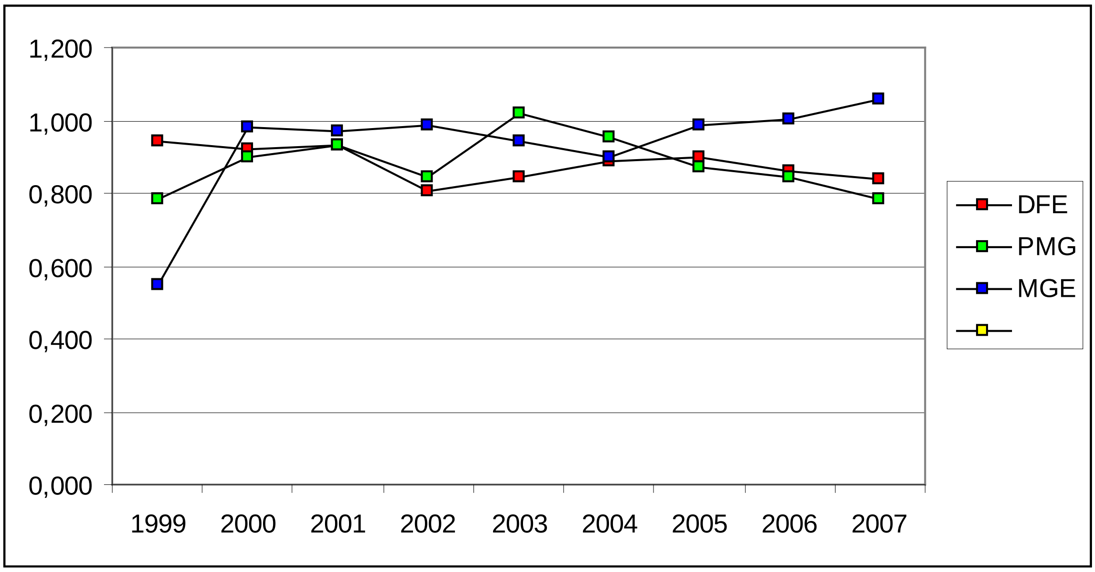
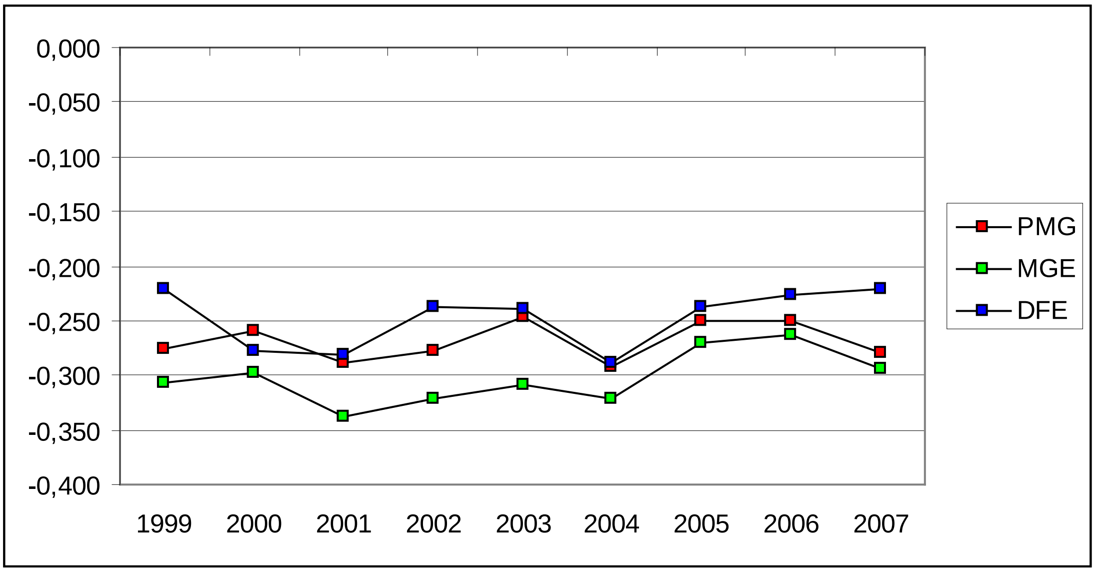


Figure 2: Saving rate coefficients for increasing sample size

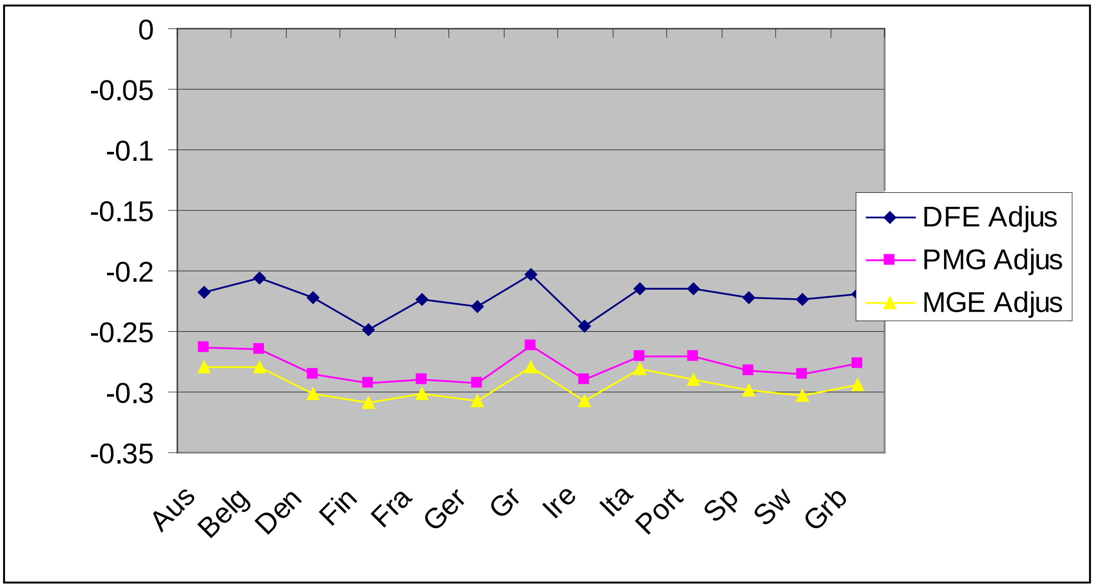


Note : The different estimators are considered as explained in the text and the sample size is extended one year at a time. Thus, the first observation shows the coefficient estimates for the sample from 1970 to 1999 (at 1999 in the chart) and the last one shows the estimates for the sample from 1970 to 2007 (at 2007).

Figure 3: Coefficients for error-correction term for increasing sample size



Note : The different estimators are considered as explained in the text and the sample size is extended one year at a time. Thus, the first observation shows the coefficient estimates for the sample from 1970 to 1999 (at 1999 in the chart) and the last one shows the estimates for the sample from 1970 to 2007 (at 2007).



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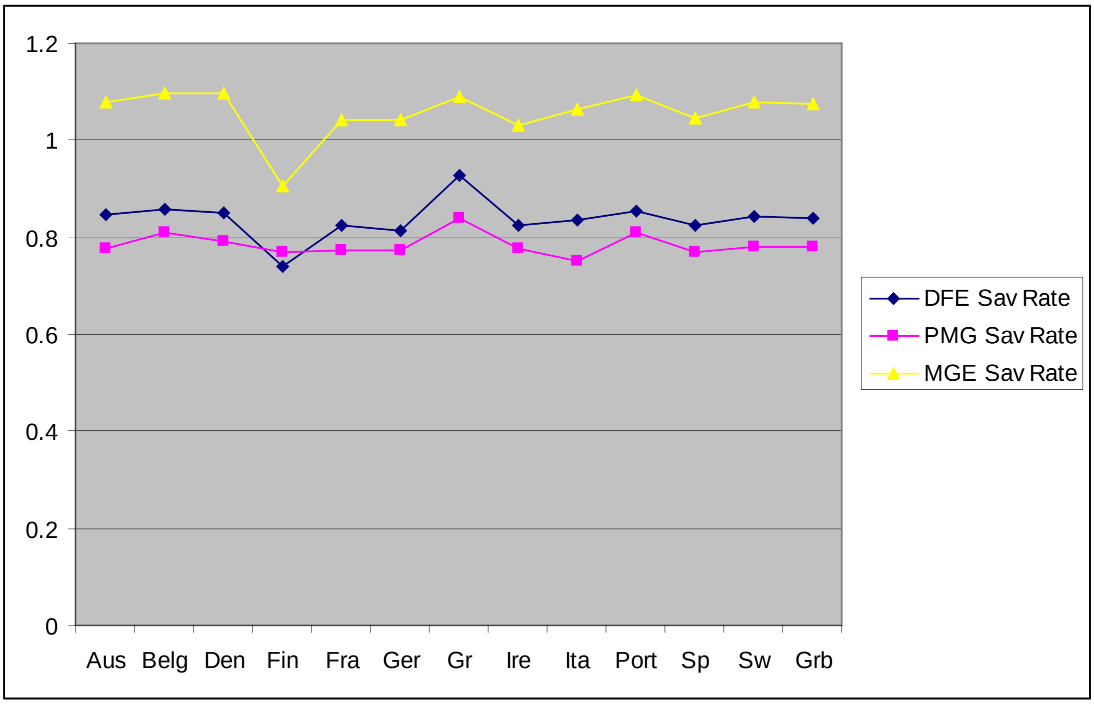
-0.1
-0.15
-0.2
-0.25
-0.3
-0.35
Aus
Belg
Den
Fin
Fra
Ger
Gr
Ire
Ita
Port
Sp
Sw
Grb
DFE Adjus
PMG Adjus
MGE Adjus



1.2
1
0.8
0.6
0.4
0.2
0
Aus
Belg
Den
Fin
Fra
Ger
Gr
Ire
Ita
Port
Sp
Sw
Grb
DFE Sav Rate
PMG Sav Rate
MGE Sav Rate