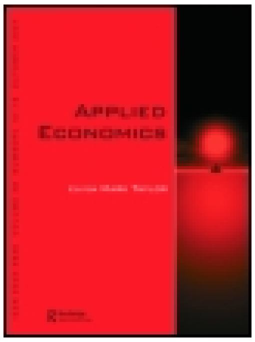

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Saving, investment and capital mobility in EU member countries: a panel data analysis of the Feldstein–Horioka puzzle

Anastassios A. Drakos[a], Georgios P. Kouretas[a,b] and Prodromos Vlamis[c,d]

[a]Department of Business Administration, Athens University of Economics and Business, Athens, Greece; [b]IPAG Business School, Paris, France; [c]Centre for Planning and Economic Research, Athens, Greece; [d]Department of Land Economy, University of Cambridge, Cambridge, UK

ABSTRACT
In this paper, we reexamine the long-standing and puzzling correlation between national saving and investment in 14 European Union (EU) countries. We employ a panel data set for the period 1970–2015 and we apply recently developed maximum likelihood panel cointegration methodologies. We find that there exists a long-run relationship between savings and investment for this panel of EU member countries, with the savings retention coefficient being low in magnitude but statistically different than zero. Therefore, we argue that there is weak evidence in favour of the Feldstein–Horioka puzzle and that the long-run international solvency condition is maintained in most of these countries. This evidence implies a moderate degree of capital mobility which is consistent with the macroeconomic experience of these countries during the period under investigation.

KEYWORDS
Capital mobility; Feldstein–Horioka puzzle; saving–investment association; panel unit root test; panel co-integration; European Union

JEL CLASSIFICATION
C22; C23; F32; F36; F41; E21

I. Introduction

International capital mobility was initially evaluated by the extent of exchange rate restrictions. However, the seminal paper of Feldstein and Horioka (1980) questioned the efficacy of this approach given the increasing evidence that capital flow takes place despite exchange rate restrictions. Eventually, two broad approaches to the evaluation of the degree of international capital mobility, namely, the price (direct) approach and the quantity (indirect) approach were broad to the surface (Obstfeld 1995). 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 other looks at actual international capital flows. The price approach is based on testing the law of one price in the context of identical financial assets. Here, the price of assets denominated in different currencies, with similar risks and maturity characteristics tends to equalize quickly through arbitrage. As Frankel (1991) points out Covered Interest Parity is used as the most appropriate indicator of the degree of financial integration and thus capital mobility. Alternatively, we could consider the quantity approach which has two main variants, the consumption smoothing approach and the saving–investment relationship.

This paper will focus on the latter approach, and in particular on what the correlation of saving and investment rates across countries may imply for the level of capital mobility. The focus on flows of capital 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 behaviour of relative rates of return have tended to concentrate on the behaviour 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.

CONTACT Georgios P. Kouretas kouretas@aueb.gr

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Interest in looking at the correlation of saving and investment across countries as a test of the degree of capital mobility stems principally from Feldstein and Horioka (1980) seminal paper. Their model is based on the simple 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 since the seminal work of Feldstein and Horioka (see, *inter alia* survey articles by Frankel 1992; Coakley, Smith, and Smith 1998; Apergis and Tsoumas 2009; references therein).

Although Feldstein and Horioka (1980) relate the presence of a low degree of correlation between savings and investment as evidence of high capital mobility later studies argued that simple saving–investment correlations may not be informative about international capital mobility (Obstfeld 1986; 1995; Baxter and Crucini 1993; Coakley, Smith, and Smith 1998 among others). Furthermore, recent studies on the saving-investment relationship have interpreted the finding of a high degree of correlation as an indication of the existence of a solvency constraint rather than low capital mobility. This approach is based on the argument that since the intertemporal budget constraint of an open economy should not allow countries to run current account deficits indefinitely (Sinn 1992), there must be a long-run relationship that ties national saving and investment together. Therefore, this approach implies that if cointegration exists between savings and investment then this provides no evidence with respect to capital mobility and it only reflects the solvency constraint (Coakley, Kulasi, and Smith 1996; 1998; Coakley and Kulasi 1997; Corbin 2004).

Furthermore within the context of cointegration Byrne, Fazio, and Fiess (2009) argue that the positive correlation between saving and investment could be explained by endogeneity if both were jointly determined by a third common force. These factors may include economic growth, population growth, productivity shocks or fiscal policies that target a balanced current account. 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. Panel data estimation techniques are appropriate to study these issues since they acknowledge cross-section-specific effects would also yield better and more representative estimation results. They propose an approach that allows us to decompose a series to its idiosyncratic level and its global level. Their main finding using data for 21 industrial countries is that rejection of the existence of cointegration between the idiosyncratic components of saving and investment and evidence in favour of the existence of a long-run relationship between their global factors is consistent with capital mobility since the long-run movements of investment are linked with long-run movements in saving at the global level.

Younas and Chakraborty (2011) argue that we should always expect some degree of home bias in the allocation of domestic savings due to barriers to entry, information cost and attitude towards risk in investing in international portfolios. They focus on the issue whether the F–H puzzle is maintained when controlling for a third factor namely the degree of openness or this will lead the savings retention coefficient to decline over time. It is argued that high savings–investment correlation can arise from excessive capital control which reduces the cross-border movement of portfolio and direct investment. By contrast, financial liberalization provides greater opportunities for domestic savings to finance investment projects that offer the highest marginal returns in the world. Therefore, empirical models that do not account for financial openness and globalization would result in an upward bias on savings retention coefficient due to the omission variables problem. Finally, Yasutomi and Horioka (2011) provide an explanation of the F–H paradox based on Adam Smith's work who argued that owners of capital invest their capital in their own country as much as they can for security reasons rather than for profit reasons.

Given the level and volatility of capital inflows experienced by many European countries during the period 1970–2015, an investigation of the validity of the F–H puzzle is highly warranted. Therefore, the present study addresses the F–H issue in the case of the EU14 countries over the period 1970–2015. In

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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. First, we analyze the F–H puzzle in the context of the joint estimation and interpretation of the short- and long-run dynamics. We follow Pelgrin and Schich (2008) and we consider the close long-run saving–investment relationship reflecting a solvency constraint and the short-run saving–investment relationship to assess the degree of capital mobility. Second, we consider the case of decomposing the savings and investment series into idiosyncratic and factor elements using the principal components methodology suggested by Bai and Ng (2004) and investigating the existence of cointegration between the idiosyncratic components as well whether factor cointegration could be identified. Finally, we control for globalization and financial openness and we estimate an augmented F–H specification using the Generalized Method of Moments (GMM) of Arellano and Bover (1995) and Blundell and Bond (1998) to account for possible endogeneity between savings and investment.

Several novel features stem from our analysis. First, based on a batter of second-generation panel unit root we are unable to reject the null hypothesis of cross-sectional dependence. Cross-sectional dependence is likely to be present in the error structure of our panel data. In fact, a large amount of literature has provided evidence on the existence of strong co-movements between macroeconomic aggregates among industrial countries (see for instance, Backus, Kehoe, and Kydland 1992). It has also been recognized that the assumption of independence across members of the panel is rather restrictive, particularly in the context of cross-country regressions. Moreover, this cross-sectional correlation may affect the finite sample properties of panel unit root tests (see O' Connel 1998). Second, given the non-stationary nature of the savings and investment ratios for the individual countries and the panel we then examine whether a long-run relationship between the two ratios exists. To this end we implemented the Johansen cointegration tests for individual countries and it was shown that for all countries a single statistical cointegration vector exists. The modified Johansen test for panel cointegration also provides strong evidence for the presence of cointegration for the panel. The evidence in favour of cointegration is further confirmed through the application of Pedroni's panel cointegration tests since the null hypothesis of no cointegration is rejected. Therefore, it is clear that the EU-14 countries maintain their inter-temporal budget solvency in the long-run. The individual counties and panel savings-retention coefficient are estimated through the Johansen methodology and the Pedroni group mean FMOLS technique with common time dummies included to control for cross-sectional dependence. For most countries the savings-retention coefficient is smaller than 0.60 and statistically significant indicating a moderate degree of capital mobility. The same evidence holds for the savings-retention coefficient for the panel. Third, we identify the existence of an association between investment and saving but this relationship is driven by global factors rather than idiosyncratic factors, which provides substantial resolution to the F–H puzzle. Finally, we show that controlling for the degree of openness with the inclusion of the KOF globalization and economic indices the saving-retention coefficient declines substantially.

The paper is organized as follows. Section II provides a theoretical background to the F–H hypothesis and reviews the methodologies used to test it as well as alternative interpretations and critiques. Section III explains the model and the econometric methodology, while Section IV describes the data and discusses the empirical results from panel data analyses whereas the final section contains the summary and conclusions.

II. Theoretical and empirical review

The Feldstein–Horioka approach: theory critique and alternative interpretations

Feldstein and Horioka (1980) proposed a simple model based on the goods market equilibrium

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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:

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

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:

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

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 Bachetta (1991), 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. Feldstein and Horioka (1980), Feldstein (1983) and Feldstein and Bachetta (1991) 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, 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, that is there is no movement of capital between countries. In order for β to be equal to 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 with cross-section data for a sample of 16 industrial countries for the period 1960–74, Feldstein and Horioka (1980) found the estimated β coefficient to be in the range 0.85-0.95 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, they appeared to be sufficient elements rigidities and preferences to keep saving invested in the country of origin.

III. Econometric methodology

The choice of the econometric model is essential to further explore the F–H puzzle and also depends on the approach that we take into consideration. Thus, for the first part of our analysis which involves the analysis of the solvency constraint argument we implement the cointegration analysis for panel data. To this end in order to test whether the savings and investment ratios are cointegrated in a heterogeneous panel of 14 European Union countries we apply two well-known panel cointegration procedures: the Johansen multivariate cointegration methodology and the Pedroni (1999, 2000, 2004) residual-

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based panel cointegration tests. Upon establishment of cointegration we then move to the estimation of the savings–retention coefficient.[1] Larsson, Lyhagen, and Lothgren (2001) have extended the standard Johansen (1991) multivariate cointegration approach used in time series analysis to accommodate for panel data heterogeneity. Larsson, Lyhagen, and Lothgren (2001) examine the mean of the standard Johansen LR (Hereafter LLL) trace test from a heterogeneous model. The LR is the simple average of the standard Johansen trace statistic for the hypothesis of reduced rank from (full rank) p to r. The hypothesis tested is

$$H_0 : rank(\Pi_i) = r_i \leq p \text{ for all } i = 1, \ldots\ldots., N$$

Against the alternative

$$H_1 : rank(\Pi_i) = p \text{ for all } i = 1, \ldots\ldots., N$$

This pair of hypotheses is the panel analogue of the time series trace statistic. In order to determine the LLL panel trace test, the statistic $Y_{\overline{LR}}$ is obtained by standardizing the average of the N countries' individual observed trace statistics, $\overline{LR}_{NT}$ as follows:

$$Y_{\overline{LR}}(H(r)/H(p)) = \frac{\sqrt{N}(L\overline{R}_{NT}(H(r)|H(p)) - E(Z_k))}{\sqrt{\mathrm{var}(Z_k)}}$$

where $E(Z_k)$ and $\mathrm{var}(Z_k)$ are the mean and variance of the asymptotic trace statistic, Z_k, and where

$$\underset{NT}{\overline{LR}}(H(r)\Big|H(p)) = \frac{1}{N}\sum_{i=1}^{N} LR_{iT}(H(r)\Big|H(p))$$

is the simple average of the standard Johansen trace statistic for the hypothesis of reduced rank from (full rank) p to r and is distributed as N(0,1). The properties of the statistic are asymptotic. Larsson, Lyhagen, and Lothgren (2001) provide the results for the trace statistic for the case of an unrestricted constant and no deterministic trends are included in the underlying VAR model. Gerdtham and Lothgren (2000) have shown that the results hold also for the case where there is a linear trend in the cointegrating space. Finally, Breitung (2005) has extended the basic model showing alternative representations of the Y_{LR} statistic by considering various deterministic components in the underlying VAR model. The Y_{LR} statistic has been demonstrated to be normally distributed asymptotically. Finally, in the panel test, the null hypothesis states that all of the N countries in the chosen panel have a maximum common rank of r cointegrating relationships among the p variables against the full rank alternative for all the countries, although each country can have its own r_i number of cointegrating vectors.

Furthermore, Pedroni (1999; 2000; 2004) panel cointegration allows for a high degree of individual heterogeneity and has no-cointegration as its null hypothesis. In order to test the null hypothesis, Pedroni (1999; 2000; 2004) developed seven cointegration statistics, comprising of four *within* dimension (panel) and three *between* dimensions (group). All are based on the residuals from the (most general) regressions

$$y_{it} = \alpha_i + \delta_i t + \beta_i \chi_{it} + \varepsilon_{it} \tag{3}$$

The first four statistics effectively pool the autoregressive coefficient in the residual based test and the other three statistics take the average, allowing more heterogeneity. Pedroni refers to the within statistics as panel cointegration statistics, and to the between statistics as group mean panel cointegration statistics, based on our discussion above. The panel tests are the panel υ-statistic (a variance bounds test), the panel ρ-statistic (analogous to the Phillips-Perron ρtest) and nonparametric and parametric panel t-statistics (or more accurately, ADF statistics). The group tests are the group ρ-statistic and the two group t-statistics. Each of these standardized distributions of the panel and group statistics are given as follows:

$$\frac{\theta_{NT} - \mu\sqrt{N}}{\sqrt{N}} \tag{4}$$

where, θ_{NT} is the corresponding observed statistic with their expected mean, μ and expected variance, ν. The one-sided test statistics are distributed asymptotic standard normal.

The second part of our analysis applies an econometric methodology that allows us to separate idiosyncratic correlation at the country level from correlation at the global level (Byrne, Fazio, and Fiess 2009). The justification for this analysis is that the positive correlation between saving and

[1]Residual based tests in single countries are well known to have non-standard distributions and low power. In panels the distribution tends to asymptotic normal, and power usually increases.

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investment could be explained by endogeneity if both were jointly determined by a third common factor. These factors include business cycle which determines both savings and investment (Obstfeld 1986; Obstfeld and Rogoff 1995) and global shocks such imported inputs and world interest rates that affect both savings and investment simultaneously (Baxter and Crucini 1993). Similar arguments have been made with respect to population growth, productivity shocks (Summer 1991; Tesar 1991; Constantini and Gutierrez 2013). Given these factors, the existence of a positive relationship between savings and investment may be uninformative about capital mobility.

We follow Byrne, Fazio, and Fiess (2009) in order to account for common shocks by using a principal components approach based on Bai and Ng (2004) and test for cointegration using Pedroni (1999, 2000, 2004) for panel data that we already employed in the first part of the analysis as suggested by Gengenbach, Palm, and Urbain (2006) and Johansen (1991) time series procedure. The investment–savings relationship can be written in panel framework as follows:

$$I_{it} = c_i + \beta_i S_{it} + e_{it} \tag{5}$$

where c_i is a set of country fixed effects and β_i are slope coefficients. If there is no long-run relationship between investment and saving and they are both unit root processes we are unable to reject the null hypothesis. Additionally, as we have already explained evidence of cointegration between the two variables is consistent with the solvency constraint and this is not informative of the degree of capital mobility.

We decompose saving and investment as follows:

$$I_{it} = c_{1i} + \lambda'_{1i} F_{1t} + \varepsilon_{it} \tag{6}$$

$$S_{it} = c_{2i} + \lambda'_{2i} F_{2t} + \eta_{it} \tag{7}$$

Following Byrne, Fazio, and Fiess (2009), by suppressing fixed effects c_{1i} and c_{2i}, we can rewrite Equation (5) as a linear combination of factors, F_{1t} and F_{2t}, and idiosyncratic components, ε_{it} and η_{it}, where λ_{1i} and λ_{2i} are factor loadings:

$$I_{it} - \beta_i S_{it} = \lambda_{1t}\left(F_{1t} - \frac{\beta_i \lambda_{2i}}{\lambda_{1i}} F_{2t}\right) + \varepsilon_{it} - \beta_i \eta_{it} \tag{8}$$

In order to identify a stationary linear combination I_{it}, $S_{it} \sim CI(1,1)$, with a cointegrating vector $(1, -\beta_i)$ two conditions need to be satisfied: $F_{1t} - (\beta_i \lambda_{2i}/\lambda_{1i})F_{2t} \sim I(0)$ and $\varepsilon_{it} - \beta_i \eta_{it} \sim I(0)$. Therefore, we require that the common factors as well as the defactored data must form a long-run relationship.

The unit root tests by Bai and Ng (2004) provide a complete procedure to test the degree of integration of a series as a first part of the analysis. In their testing framework the common factors, instead of being treated as a nuisance, become a direct object of further investigation. In contrast to Pesaran (2007) or Moon and Perron (2004) the PANIC (Panel Analysis of Nonstationarity in Idiosyncratic and Common components) model permits the nonstationarity in a panel of observed data to come either from a common source, or from the idiosyncratic errors, or both.[2] Therefore, they focus on consistent estimation of the common factors and error terms, to test the properties of these series separately. They decompose a series y_{it} as a sum of a deterministic part, a common component expressed as a factor structure and an error that is largely idiosyncratic. Then the process y_{it} is nonstationary if one or more of the common factors are non-stationary, or the idiosyncratic error is non-stationary, or both. Instead of testing for the presence of a unit root directly in y_{it}, Bai and Ng (2004) propose to test the common factors and the idiosyncratic components separately. This is the main difference with respect to other testing procedures based on factor structure, which generally test the unit root only in the de-factored data. For that, they have to use a decomposition method of the data which is robust to the degree of integration of the common or idiosyncratic components. In other words, the common variations must be extracted without appealing to stationarity assumptions and/or cointegration restrictions. Bai and Ng (2004) accomplish this by estimating factors on first-differenced data and cumulating these estimated factors.

Following the identication of the idiosyncratic and common factor components we then test for cointegration between the idiosyncratic components, $\varepsilon_{it} - \beta_i \eta_{it} \sim I(0)$, using Pedroni (2004) tests for panel cointegration and between the factors

[2]In Pesaran's (2007) or Moon and Perron's (2004) model the data under the unit root null contain a common, as well as an idiosyncratic stochastic trend.

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$F_{1t} - (\beta_i \lambda_{2i}/\lambda_{1i})F_{2t} \sim I(0)$, by applying the standard Johansen (1991) multivariate (1991) approach for time series. It is then argued that no evidence of cointegration between the idiosyncratic components will provide a possible resolution of the F–H puzzle, whereas evidence of cointegration between factors would imply that industrial countries have a global factor in investment, which is associated with a global factor in saving.

The final part of our analysis follows Corbin (2001) and Younas and Chakraborty (2011) in which we control for country-specific effects. Corbin (2001) argues that the high savings retention coefficient could not be considered as evidence of low capital mobility but is due to the existence of country-specific effects. Therefore, we estimate the savings–investment relationship taking into consideration that: (a) endogeneity of domestic savings that may be present; (b) the dynamic relationship between domestic savings and investment as both are impacted by the prior values of each other; and (iii) the existence of unobserved country-specific effects. For these reasons we employ Arellano and Bover (1995) and Blundell and Bond (1998) GMM methodology for dynamic panels. Besides accounting for the specified dynamics, the Blundell-Bond estimator has two additional virtues. First, it does not break down in the presence of unit roots (for a proof see Binder, Hsiao, and Pesaran 2003). Second, and most important, it accommodates the possible endogeneity between the investment and some of the right-hand side variables by means of appropriate instruments. In addition we take one year lagged values of all independent variables in our model. The lagged value of the savings is justified by the fact that it takes time for savings to be transformed into fixed capital formation, which is the typical measure of investment. Therefore we estimate the following dynamic augmented F–H specification in first differences:

$$\left(\frac{I}{Y}\right)_{it} = \alpha_0 + \alpha_1 \left(\frac{I}{Y}\right)_{i,t-1} + \alpha_2 \left(\frac{S}{Y}\right)_{i,t-1} + \alpha_3 (X)_{i,t-1} + \alpha_4 X \left(\frac{S}{Y}\right)_{i,t-1} + u_{it} \quad (9)$$

where subscripts i and t indicate country and time period, respectively.

In Equation (9) we include the interaction effect between the globalization index and savings as a percentage of GDP $X\left(\frac{S}{Y}\right)$ in order to capture the potential impact of increasing overall and economic globalization over time on the degree of capital mobility. A problem with the inclusion of interaction effects is the severe multicollinearity between the multiplicative term and its constituents that may be present. Therefore, we apply a procedure which consists of transforming the values of these variables to deviations from their means, and then forming the product term from these deviations.

IV. Empirical analysis and findings

Data for gross domestic saving and gross fixed capital formation (investment) as a percentage of GDP for 14 European Union countries members have been collected from *Datastream*. In particular, we have selected a sample of annual data from 1970 to 2015, allowing us a panel of $N = 14$, $T = 46$ dimensions, a total of 644 annual investment-savings data. The list of countries includes Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, Spain, Sweden and the United Kingdom. Both variables are measured in millions of euro and current prices. The choice of countries and of the period is dictated by the need to construct a consistent balance data panel for the empirical analysis. Furthermore, compared to most previous studies our sample includes the period of the recent financial crisis 2007–2009 as well as the ongoing Eurozone debt crisis.

We also use the KOF globalization index of Dreher (2006) as a formal measure of economic openness and financial markets liberalization. The KOF globalization index is a weighted average of economic, political and social globalization. Furthermore, we also use economic globalization as a weighted index of actual economic flows (both trade and capital flows), and the index of restrictions on trade and capital flows. The data for both indices are available in http://globalization.kof.ethz.ch/. Both indices take a value in the range 0–100. Higher value of a globalization index corresponds to more openness. The mean value of globalization index equals 74.28 with a standard deviation of 13.89 whereas the mean value for the economic globalization index equals 68.21 with a standard deviation of 14.28.

Table 1 reports the results from the application of several second-generation panel unit root tests in

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Table 1. Second generation panel unit root tests.

	Moon and Perron (2004)			Choi (2002)			Chang (2002)	Breitung and Das (2005)	Harris, Leybourne, and McCabe (2005)	Phillips and Sul (2003)
Variables	t_a^*	t_b^*	$\hat{\rho}_{pool}$	P_m	Z	L^*	S_N	t_{rob}	$\hat{S}_k^F$	Z_{PS}
(I/Y)	−6.629	−14.938	0.815	2.392	−6.331	−3.559	2.7466	−0.6298	2.719	31.065
	(0.973)	(0.000)		(0.144)	(0.225)	(0.089)	(0.985)	(0.264)	(0.004)	(0.185)
(S/Y)	−5.265	−11.625	0.857	2.319	−3.412	−3.467	3.6942	−1.1802	2.322	26.077
	(0.418)	(0.000)		(0.079)	(0.262)	(0.239)	(0.988)	(0.121)	(0.005)	(0.365)

1. t_a^* and t_b^* are the unit root test statistics based on de-factored panel data (Moon and Perron 2004).
Corresponding p-values are in parentheses; $\hat{\rho}_{pool}$ is the corrected pooled estimates of the autoregressive parameter.
t_a^*,t_b^* and $\hat{\rho}_{pool}$ are computed with a Quadratic Spectral kernel function in spite of a Bartlett kernel function, and the bandwidth parameters are computed according to the Newey and West (1994) procedure.
2. The P_m test is a modified Fisher's inverse chi-square test (Choi 2001). The Z test is an inverse normal test.
The L^* test is a modified logit test. All statistics have a standard normal distribution under H_0 when T and N tend to infinity (Choi 2002). The null hypothesis of non stationarity is rejected when Pm is greater than the upper tail of the standard normal distribution. For other tests, the null is rejected when the realizations is inferior to the lower tail of the standard normal distribution. Corresponding p-values are in parentheses.
3. The S_N statistic corresponds to the average of individual nonlinear IV t-ratio statistics (Chang 2002). It has a N(0, 1) distribution under H_0. The Instrument Generating Function (IGF) is: IGF 2: F(x) = I(|x|<K). Trimmed OLS on [-K,K] where K is the second quantile of $y_{i,t}$
4. t_{rob} is the robust t-statistic free of size distortions due to contemporaneous cross-sectional correlations for N and $T \rightarrow \infty$(Breitung and Das 2005).
5. Z_{PS} is the Phillips and Sul (2003) test for cross section dependence being caused by a common factor. Computational work was performed in GAUSS. A GAUSS code is available from Donggyu Sul. Here, the lag orders of the univariate ADF regressions are chosen based on the top-down method. The maximum number of lags was set to 10.
6. $\hat{S}_k^F$ statistic is a panel stationarity test (Harris, Leybourne, and McCabe (2005). Following Harris, Leybourne, and McCabe (2005) we implement $\hat{S}_k^F$ statistic using $k = \left\lceil (3T)^{1/2} \right\rceil$ and the long run variances are estimated using a Bartlett lag window with $l = \left\lceil 12(T/100)^{1/4} \right\rceil$.Following Bai and Ng (2004) the long run variances are computed using the Quadratic Spectral lag window $l = \left\lceil 12(T/100)^{1/4} \right\rceil$ giving $\bar{\eta}_\mu^{QS}$.

order to examine the presence of cross-section dependence. First, we apply the Phillips and Sul (2003) factor structure approach to examine dependence across units. The inverse normal Z statistic strongly implies that the unit root hypothesis should not be rejected for almost all individual series. This meta-statistic combines the p-values of the individual ADF regressions of the de-factored data. We then provide the results of the panel stationarity test of Harris, Leybourne, and McCabe (2005) that allows for at least one factor which also strongly reject the null stationarity in our panel data with the $\hat{S}_k^F$ statistic.

These first three tests reject the unit root hypothesis when allowing a single factor structure in the composite error term. Table 1 also reports the results of the Moon and Perron (2004) test which requires multiple common factors and thus it evaluated whether the one-factor model is sufficient modelling framework of the dependence structure of the panel data. A set of information criteria to the saving and investment rates in the balanced EU 14 countries panel were applied and congruent results obtained when setting $k^{\max} = 6$ and focusing on BIC_3 in which case it is recommended to assume one common factor.[3] The long-run variances were estimated using the Andrews and Monahan (1992) method. The evidence we obtain for the application of the two test statistics is inconclusive. The t_α^* statistic implies that the null of a homogeneous unit root in the panel for the assumption of one, two or six common factors should not be rejected. In contrast, the t_b^* statistic rejected the null for all specifications as indicated in Table 1.[4] For robustness purposes, Table 1 reports evidence from three additional tests for nonstationarity. First, Choi (2002) test which allows for cross-sectional dependence and has a specification based on error component model and the results are in favour of the presence of a unit root. Second, the tests of Breitung and Das (2005) and Chang (2002) clearly fail to reject the joint nonstationarity hypothesis.[5]

Based on the evidence provided by the panel unit root tests that the investment and saving rate series are non-stationary in levels, we pursue further analysis on the possible evidence of cointegration in a more rigorous setting. Table 2 reports our results from the application of panel and group statistics proposed by Pedroni (1999; 2000; 2004) and we conclude that we reject the null of no cointegration. Therefore, the establishment of

[3]The latter is the preferred criterion of Moon and Perron in small samples.
[4]However, none of the simulations studies in Moon and Perron test (see Gengenbach et al., 2006; Gutierrez 2006) provides guidance as to which t statistic should be preferred in its applications.
[5]The results of the unit root and stationarity tests clearly show that our cointegration results are robust to the choice of the unit root test.

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Table 2. Pedroni's Heterogeneous panel cointegration tests.

Test Statistic	Statistic	Probability
Within-dimension		
Variance ratio statistic	2.334	0.0058
Rho statistic	−3.121	0.0015
Phillips-Perron statistic	−3.229	0.0006
ADF *t*-statistic	−4.339	0.0000
Between-dimension		
Rho statistic	−3.665	0.0005
Phillips-Perron statistic	−2.457	0.0077
ADF *t*-statistic	−3.688	0.0000

Heterogeneous 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 and Perron (1988) rho-statistic. The third is a panel analogue of the Phillips and 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.

a long-run relationship between saving and investment rate in the panel of the economies are economically meaningful since it supports the argument that these countries fulfill the long-run solvency constraint.

To provide robustness to these results we also apply the Johansen multivariate cointegration approach for panels discussed in Section 3. The results are reported in Table 4. The estimate LR trace test statistics have been adjusted for small sample as suggested by Reimers (1992). For individual countries, the number of lags used in the VAR model is equal to four and in addition the estimated model is assumed to have a linear deterministic trend.[6] For the case of individual country, we find strong evidence for the support of one cointegrating vector for all countries since the null hypothesis $r = 0$ is rejected whereas we were unable to reject the null hypothesis of $r = 1$ for all cases. The results for the individual countries are consistent with the long-run solvency of these countries' inter-temporal budget constraint. For the panel cointegration rank test, the observed Z_{LR} statistic of 16.82 is greater than the critical $Z$of 1.64 and therefore we reject the null hypothesis of no cointegration against the evidence of a statistically significant cointegration vector whereas the value of the Z_{LR} statistic of 2.15 for the null hypothesis of one cointegration vector against two cointegrating vectors could not be rejected. Therefore the evidence from the LLL panel test shows that (I/Y) and (S/Y) are cointegrated. This finding implies that for the panel as a whole, the inter-temporal solvency condition that (I/Y) and (S/Y) is not violated as their gap cannot deviate permanently.[7]

Given the overwhelming evidence that a cointegrating relationship and one cointegrating vector exists between the two variables, (I/Y) and (S/Y), we now move to the estimation of the long-run savings coefficient β using a panel cointegration estimator. The estimated coefficients are obtained from the Johansen procedure and the Pedroni (2001; 2004) group mean fully modified ordinary least squares (GMFMOLS) which has several advantages over alternative estimation approaches. In Tables 3 and 4, we report the estimated long-run savings–retention coefficients (Feldstein–Horioka coefficient) for both the individual countries and the panel from both estimation techniques. To capture possible cross-sectional dependence in the data sample, common time dummies are included in the estimated model. Our results show that the esti-

Table 3. GMFMOLS: Individual and panel saving–investment retention coefficients.

Countries	FMOLS	*t*-statistic
Austria	0.767*	5.532
Belgium	0.423*	3.167
Denmark	0.494*	2.922
Finland	0.871*	0.312
France	0.538*	4.355
Germany	0.519*	3.559
Greece	0.722*	2.776
Ireland	0.466*	2.831
Italy	0.755*	5.989
Netherlands	0.718*	3.445
Portugal	0.338*	4.339
Spain	0.709*	3.809
Sweden	0.437*	3.199
UK	0.761*	3.672
Panel Results		
(Without common time dummies	0.651*	6.335
Wald test: $\beta = 1$	144.09*	10.554
(With common time dummies)	0.578*	11.743
Wald test: $\beta = 1$	221.65*	13.221

Number in brackets are *t*-statistics.
*Denotes statistical significance at the 5% critical value respectively.

[6]We obtained similar results when we consider different lag structure and alternative specifications regarding the inclusion of a constant and a constant and linear trend.

[7]The intuition behind the use of the modified Johansen test for panels is identical to the original Johansen test for time series data that is to reveal the existence of a long-run relationship which is the solvency constraint (see last paragraph p. 9). We use both panel and time-series Johansen test since it has been shown (Karlsson and Lothgren 2000) that while using panel enhances the power of unit root tests, these tests may actually lack power when a large fraction of the data is stationary. Therefore, they recommend a careful joint analysis of both the individual country unit root test for each cross-sectional unit and the panel unit root tests when the stationary properties of the panel series need to be evaluated.

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Table 4. Likelihood-based cointegration tests: panel and individual country test results.

	$LR_{iT}(H(r)\|H(2))$			
Country	$r = 0$	$r = 1$	β	Rank(r_i)
Austria	25.56	4.80	0.776*	1
Belgium	30.62	8.25	0.421*	1
Denmark	39.52	11.40	0.416	1
Finland	26.30	6.56	0.889	1
France	16.24	6.03	1.141*	0
Germany	18.17	6.37	0.998*	0
Greece	38.92	7.62	0.738*	1
Ireland	27.78	9.91	0.065	1
Italy	28.60	8.18	1.015*	1
Netherlands	28.57	10.19	0.713*	1
Portugal	30.49	9.49	0.390*	1
Spain	10.45	2.85	0.698*	0
Sweden	18.64	6.34	0.451	0
UK	33.35	5.55	0.577*	1
5% critical value	25.87	12.52		
Panel Tests				
LR_{NT}	25.338	6.228		
$E(Z_k)Var(Z_k)$	15.31	6.16		
	23.39	9.33		
$Z_{\overline{LR}}$	6.82	2.15		
N	14	14		

We estimate a model with a deterministic trend in the cointegration space. Asymptotic moments for standardizing the test statistic are from Breitung (2005). For individual country tests, the 5% critical values for testing $r = 0$ and $r = 1$, respectively are 25.87 and 12.52.
*Denotes significance at the 5% level of significance. The critical value for the panel test is 1.645.

mated savings–retention coefficients have the expected sign for all individual countries and they are all statistically significant. Furthermore, we underline that for most cases the magnitude of the coefficient is smaller than the cut-off value of 0.60 suggested by Murphy (1984) indicating that these countries not only have maintained the long-run solvency constraint, although at the same time these coefficients indicate that there is some degree of capital mobility but not perfect capital mobility as argued by Feldstein and Horioka (1980). Moreover we obtained similar evidence from the estimated savings–retention coefficient for the panel as a whole since their values are relatively small and statistically different from either zero or unity at the 5% level of significance according to a standard Wald test. Therefore, we conclude that the intertemporal solvency constraint is confirmed for the EU-14 countries but some capital mobility is also observed.

We now turn to the analysis of the factor structure with the application of the Bai and Ng (2004) by decomposing the saving and investment variables 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 non-stationary either the idiosyncratic error or some of the common factors should be non-stationary. 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 $ADF^c_{\hat{F}}$ 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 (2002). Results of Bai and Ng (2004) test, reported in Table 5, are in favor of unit root for our variables. Idiosyncratic shocks to each variable are all non-stationary. 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, that is at least two independent non-stationary common factors can be identified.

Given that the idiosyncratic components of saving and investment are non-stationary we apply the panel cointegration test developed by Pedroni (2004). Based on the results given in Table 6 we conclude that we are unable to reject the null hypothesis of no cointegration between the idiosyncratic component of saving and

Table 5. Results of Bai and Ng (2004) unit root tests.

					Common factors	
		Idiosyncratic shocks			Trends $\hat{r}_1$	
	$\hat{r}$	$Z^c_{\hat{e}}$	$P^c_{\hat{e}}$	$ADF^c_{\hat{F}}$	MQ_c	MQ_f
(I/Y)	1	1.0128	34.3314	−1.698	-	-
		(0.1678)	(0.102)	(0.333)		
(S/Y)	3	1.2029	35.3928	-	3	3
		(0.1945)	(0.1129)			

$\hat{r}$ is the estimated number of common factors, based on BIC criterion function. For the idiosyncratic components $\hat{e}_{it}$, only pooled unit root test statistics are reported.
$P^c_{\hat{e}}$ is a Fisher s type statistic based on p-values of the individual ADF tests. Under H_0, $P^c_{\hat{e}}$ has a $x^2(2N)$ distribution when T tends to infinity and N is fixed.
$Z^c_{\hat{e}}$ is a standardized Choi's type statistic for large N samples: under H_0, $Z^c_{\hat{e}}$ has a N (0, 1) distribution; p-values are in parentheses. For the idiosyncratic components $\hat{F}_t$, two cases must be distinguished: if $\hat{r} = 1$; only standard ADF t-statistic, denoted $ADF^c_{\hat{F}}$ is reported with its p-value. If $\hat{r} > 1$, the estimated number $\hat{r}_1$ of independent stochastic trends in the common factors is reported (columns 6–7). The first estimated value $\hat{r}_1$ is derived from the filtered test MQ_f and the second one is derived from the corrected test MQ_c. The significance level of these tests is 5%.

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Table 6. Idiosyncratic and factor cointegration.

PANEL A. Pedroni (2004) *idiosyncratic cointegration*		
	Panel	Group
Rho statistic	−1.677	−0.225
Phillips-Perron statistic	−1.599	−0.377
ADF *t*-statistic	−1.622	−1.449
PANEL B. Johansen (1991) *factor cointegration*		
r	Trace test	5% critical value
0	22.15*	20.26
1	3.55	9.16

*Denotes statistical significance at the 5% critical level. The panel tests statistics are distributed as a standard normal and have a critical value of −1.64 at the 5% level. The 5% critical values for the Trace test are taken from MacKinnon, Haug, and Michelis (1999; Table III).

investment at the 5 per cent level of significance. The finding of no significant long-run relationship between could be interpreted as evidence of perfect capital mobility and this may provide a resolution of the Feldstein–Horioka puzzle. Decomposition also implies that the saving and investment principal components are across countries but not across time and therefore we use standard time series cointegration techniques (Johansen 1991; Johansen and Juselius 1992) to test for the existence of a long-run relationship between these factors. Based on the standard trace and maximum eigenvalue likelihood ratio tests we reject the null hypothesis for no cointegration and we are able to identify one statistically significant cointegration vector. In line with Byrne, Fazio, and Fiess (2009) this finding implies that there is an association between investment and saving but this relationship is driven by global factors.

Table 7 presents our estimates from the GMM estimates for the F–H equation without the inclusion of globalization and its interaction term with the savings rate. Given the availability of the KOF globalization indices are available until 2013 and therefore the GMM estimates are done for the period 1970–2013. In regression I we include lagged value of the dependent variable to account for possible persistence. The results show that the savings retention coefficient is positive and significant at the 5 percent level for the sample of the EU-14 countries. However, the size of the coefficient is statistically different from unity and therefore the argument in favour of home bias exhibit by investors has substantial merit.

We now turn our focus on the key variable of globalization. In regression II it is shown that the coefficient of globalization is positive and statistically significant, which implies that an increase in overall globalization leads to an increase on investment. In addition this result suggests that countries with more economic openness can run larger current account deficit due to easier access to international financial markets. Furthermore, the interaction effect of globalization and savings rate is negative and statistically significant. This finding is consistent with the theoretical argument that with increased economic and financial integration, which is the case in the EU region, capital often flows to investment projects that offer the highest marginal returns and consequently leads to a decline in the coefficient of savings.

Finally, in regression III we control for economic globalization which as we have already explained measures in a straightforward way financial and trade flows as well as the restrictions on trade and capital movements. It is clear from the GMM estimates that economic globalization has a larger effect on the investment than when we use the overall globalization index.[8] Furthermore, it is important again to note that the coefficient of the interaction term is negative and statistically significant. Therefore, this implies that for a rise of economic globalization by 10 units will result to a decline in savings by 0.10 units. These findings suggest that as economic and financial integration increases within the EU-14 region home bias in the allocation of domestic savings decreases. As financial market liberalization increase and countries lower restrictions on capital movements, it is clear that domestic investment not only depends on domestic savings but is

[8]We observe in both regressions I and II the savings retention coefficient (main effect) has a much higher values as compared to the one obtained without controlling for globalization and economic globalization. This is a standard outcome in these IV and GMM estimation methods.

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Table 7. Interest and savings rates: dynamic panel regressions.

	I	II	III
Lagged Investment Rate	0.313	0.186	0.148
	(5.81*)	(2.59*)	(3.03*)
Lagged Saving Rate	0.484	0.799	0.978
	(10.61*)	(7.23*)	(9.01*)
Lagged Globalization		0.334	
		(8.12*)	
Lagged (Globalization* Savings Rate)		−0.007	
		−0.007	
Lagged Economic Globalization			0.277
			(8.03*)
Lagged (Economic Globalization* Savings Rate)			−0.011
			(6.27*)
Descriptive statistics			
Obs	616	616	616
Wald-test	198.66	403.23	639.22
Serial correlation (*p*-value)	0.588	0.686	0.792
Sargan (*p* -value)	0.129	0.256	0.308

The table reports coefficients and *t*-statistics (in parentheses). In all regressions dependent variable is the investment ratio. The explanatory variables are as follows: lagged investment ratio, lagged savings ratio. obs is the number of observations, the Wald-test and its associated *p*-value denote the goodness of fit of the regressions, Serial correlation is the test for second order autocorrelation and Sargan is the test for overidentifying restrictions. *,**and ***Denote statistical significance at the 5% level, respectively.

financed by the pool of global savings. Therefore, given the significance of the degree of openness we argue that there is support in favour of a weaker correlation between domestic savings and domestic investment which gives substantial support of the F–H argument of a higher degree of capital mobility.[9]

V. Summary and concluding remarks

In this paper we re-examine the validity of the Feldstein–Horioka hypothesis that low domestic saving-investment correlation implies high degree of international capital mobility for a heterogeneous panel for 14 European Union countries members using annual data over the period 1970–2015. The recent financial crisis and the Eurozone debt crisis are included in our empirical investigation. We analyze the F–H puzzle focusing on three different but at the same time complementary approaches given the non-stationary nature of the data.

Thus, we documented the EU-14 countries maintain their inter-temporal budget solvency in the long-run. The individual country and panel savings–retention coefficient were shown that for most countries the savings–retention coefficient is smaller than 0.60 and statistically significant indicating a moderate degree of capital mobility. The same evidence holds for the savings–retention coefficient for the panel. These findings are in line with studies like those of Coakley and Kulasi (1997), Corbin (2004) and Pelgrin and Schich (2008).

Furthermore, we decomposed the two variables in their idiosyncrating and factor components using a variant of principal components analysis developed by Bai and Ng (2004). We conclude that there is no relationship between the idiosyncratic pasts of savings and investment, a finding which is consistent with the main argument of F–H and against the argument of inter-temporal budget solvency given that the only association between the two variables found is globally determined. Our results further confirm the results of Byrne, Fazio, and Fiess (2009), Younas and Chakraborty (2011) and Constantini and Gutierrez (2013) for the importance of global factors in explain the F–H puzzle.

Finally, capital account liberalization and the integration of world financial markets should increase capital mobility across countries. In the last part of our analysis we control for the degree of openness and thus we examine the impact of globalization of the savings–investment relationship. Using the GMM method we find substantial evidence that globalization has led to an increase in the degree of capital mobility over time. We also argue that countries with more financial openness can run higher current account deficits due to access to foreign capital markets.

The overall results may lead to the conclusion that the Feldstein–Horioka puzzle is partially valid for the panel of EU-14 countries. These empirical

[9]A referee pointed out that the analysis should also discuss for robustness purposes the possible significance of the recent financial crisis on the savings-investment relationship. On this regard we note the following. First, the possible impact of the financial crisis 2007–2009 and/or the Eurozone debt crisis 2009-present will be very limited because the data is annual and the savings and investment variables do not behave like exchange rates and/or other 'jump' variables in which case detection of these type of effects will be more pronounced. What is expected to find is that during turbulence periods like the current period home has increased given that risk premia also has risen and that makes the savings-investment relationship stronger with the saving retention coefficient to have risen. Therefore, the two variables became more linked for Eurozone countries and the European Union countries during the financial crisis and during the early stages of the debt crisis. This is quite possible given that in the peak of the financial crisis and for several months afterwards the interbank money market was not really functioning and international capital flows diminished during this time. As time moves forward and the status of the euro evolves, whether this temporary shock was in fact temporary, as appears to be the case, can and should be continue to be assessed.

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findings are related with the ongoing financial integration within EU over the last two decades although the fact that these coefficients are still statistically different than zero may also reflect the fact that the financial integration has not been completed and the banking union is still in its early stages.

Disclosure statement

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