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Re-examining the Feldstein–Horioka and Sachs' views of capital mobility: A heterogeneous panel setup

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ABSTRACT

We re-examine two complementary views of international capital mobility using data for 25 OECD countries over the period 1970–2011. Estimation of the original Feldstein–Horioka and Sachs' equations provides mixed evidence of capital mobility, though we do not detect a significant bias towards finding in favour of capital immobility in using time-averaged data. However, potential bias in cross-sectional estimation motivates us to examine the data as a heterogeneous panel which allows us to control for the effects of cross-sectional dependence and endogeneity. In addressing the Feldstein–Horioka puzzle, application of the CCEMG estimator of Pesaran (2006) to the Feldstein–Horioka and Sachs' equations points towards greater (though not perfect) capital mobility than hitherto found.

1. Introduction

Recent events surrounding the global financial crisis and its aftermath have highlighted the profound macro- and microeconomic implications of high international capital mobility. Measuring capital mobility, however, has proved to be problematic. On the one hand, one might follow Frankel (1992) and others by using covered interest parity as an appropriate indicator of the degree of financial integration and therefore capital mobility across national boundaries. An alternative way forward is to consider a more indirect approach that concentrates on the effects of capital mobility on macroeconomic aggregates such as the relationship between domestic savings and investment (Feldstein & Horioka, 1980). Here it is argued that in a world with a perfect mobility of capital, domestic savings would search for the highest returns in the world capital markets independent of domestic investment demand and, in the same way, world capital markets would satisfy domestic investment needs independent of the supply of local savings (Taylor, 1996). Domestic savings would react to the international rates of return and so investment would be funded from the world capital market through a current account deficit. If, however, capital were perfectly immobile, then one would expect domestic savings and investment to be characterised by a correlation coefficient of unity.[1]

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[1] Commencing with the work of Obstfeld (1994), it has been argued that conditional on consumers exploiting all risk-sharing opportunities, financial integration produces benefits in terms of consumption risk sharing and smoothing, as domestic consumption growth becomes less correlated with domestic income growth, but more correlated with world consumption growth. See, for example, Suzuki (2014), who provides recent empirical estimates of the extent to which domestic consumption and income are delinked because of financial integration.

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Obtaining a correlation of savings and investment close to one in their cross-section analyses for industrialised OECD countries for the 1960–1974 period, led Feldstein and Horioka (1980) to reject the capital mobility assumption. The presence of a high correlation between domestic savings and investment despite the easing of capital controls has constituted a puzzle that has been the focus of an extensive literature. Such studies have employed cross-section and/or time series data in an attempt to reconcile Feldstein and Horioka's (F–H) results with the capital mobility hypothesis; see, *inter alia*, survey articles by authors such as Lapp (1996), Coakley, Smith, and Smith (1998) and Apergis and Tsoumas (2009) and references therein. Following Rao, Tamazian, and Kumar (2010), Ketenci (2013) and others, it seems reasonable to draw on the findings of a limited number of key studies drawn from a large literature. The original F–H findings were initially confirmed by Feldstein (1983) and Feldstein and Bacchetta (1991), who extended the sample period to include the post-Bretton Woods agreement. Among the studies using cross-sectional or panel methods focussed on OECD countries, Tesar (1991) estimates the *savings retention coefficient* or proportion of incremental savings that is invested domestically to be of the order of 0.85, and Coakley, Fuertes, and Spagnolo (2001) confirm that capital mobility exists in the long-run at least with a savings retention coefficient of 0.32. Further work by Giannone and Lenza (2010) employing a Factor Augmented Panel Regression technique sees a slightly higher savings retention coefficient of the order of 0.35 over various sub-samples embodied by a 1970–2007 study period, while Kollias, Mylonidis, and Paleologou (2008) find much lower panel-based regression coefficients of around 0.15. Katsimi and Moutos (2009) apply OLS methods to broader definitions of savings and investment (that account for the accumulation of human capital) and obtain a savings retention coefficient which is about 0.5. Fouquau, Hurlin, and Rabaud (2009) use a Panel Smooth Threshold Regression Model and provide estimates of the savings retention coefficient of around 0.65 for the period 1960–2000. Di Iorio and Fachin (2007) utilise panel tests allowing for cointegration between savings and investment in the long run with a structural break. Their country specific estimates of the savings retention coefficient range from 0.59 to 1.03 based on a fully modified OLS for the study period 1960–2002. Lastly, a recent evidence by Ketenci (2013) employing panel dynamic OLS estimates a savings retention coefficient of 0.27 in the case of the OECD over the study period 1970–2008.

In this paper, we approach the F–H puzzle by taking the view that it can be adequately addressed through a heterogeneous panel estimation that accommodates the presence of cross-sectional dependence. By doing this, we are able to provide a more realistic assessment of the extent of capital mobility which turns out to be greater than hitherto thought. It seems reasonable to argue that countries will differ in the correlation between domestic savings and investment. Furthermore, it is plausible that savings and investment across counties will have a tendency to move together over time. The relationship between domestic savings and investment may be driven by common factors such as productivity shocks. The failure to take on board these latter considerations can give rise to size distortion and erroneous inference drawn from the estimated domestic savings–investment correlations. In order to address these important economic considerations, we examine the F–H hypothesis through the application of estimators from the heterogeneous panel literature, including the Pesaran (2006) cross correlated effects mean group (CCEMG) estimator. A key advantage associated with the CCEMG estimator is that in assessing the relationship between domestic savings and investment, endogeneity can be accommodated when it arises from the common factors driving both the dependent and independent variables. Coakley, Fuertes, and Spagnolo (2004) first apply this estimator to examine the F–H puzzle using a panel data set of 12 OECD countries over 1980Q1–2001Q4. Payne and Kumazawa (2006) then extend the analysis by Coakley et al. (2004) to a panel of 47 developing countries over the period 1980–2003. Their estimates of the slope coefficient in the investment–savings equations are lower than those based on traditional cross-sectional estimations, indicating higher capital mobility.

We contribute to the previous literature by using data on 25 OECD countries over the period 1970–2011, and applying the CCEMG estimator not only to the F–H equation, but also we provide a first application of this estimator to the Sachs (1981) equations, which consider the relationship between the current account and savings and investment ratios. This latter approach takes into account the role of savings and investment in current account dynamics. Indeed, if the complete absence of capital mobility means that domestic savings and investments move closely in tandem, then the current account should be unaffected by fluctuations in domestic savings or investments. We believe that the simultaneous examination of the empirical approaches of F–H and Sachs provide a more complete view of the international mobility of capital. The former is based on the consideration of two domestic variables for a given country, whereas the latter involves analysing the behaviour of the external deficit of that country with respect to the rest of the world. The choice of the sample period can be justified on the grounds that the previous four decades have seen on major institutional changes that have considerably eased the mobility of capital, especially among developed countries. This should be reflected in a lower correlation between domestic savings and investment. Using our sample of 25 OECD countries, this is indeed the case when we conduct a year-by-year estimation of the investment–savings relationship as opposed to estimation using time-averaged data. In contrast to the previous cross-sectional estimates, we do not detect a significant upward bias in the savings and investment correlation involving time-averaged data. In turn, the Sachs equations point to the insignificance of investment in driving the current account. This implies that investment and savings are correlated, which is in turn consistent with the absence of capital mobility. However, when examining the data within a heterogeneous panel framework with cross-sectional dependence, both savings and investment turn out to be factors that help explain the behaviour of the current account pointing to stronger capital mobility.

The structure of this paper is as follows. Section 2 describes the data used in the analysis. Section 3 presents the results of the empirical analysis and discusses their economic implications. Section 4 concludes.

2. Data

The data set, obtained from the OECD iLibrary (as downloaded on 23 January 2013), consists of 42 annual observations from 1970 to 2011 on investment, gross savings and the current account (where the latter is calculated as the difference between gross savings and investment) for 25 OECD member countries. The sample of countries includes Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, Turkey, United Kingdom and United States. This is an extended group of countries compared to those originally analysed by Feldstein and Horioka (1980) and Feldstein (1983).[2] All three variables (namely investment, gross savings and the current account) are measured in levels in US$ millions, current prices and current PPPs, and for the purposes of the empirical analysis are expressed as a proportion of GDP. The choice of countries and of the sample period is dictated by the need to assemble a consistent balanced data panel for the purposes of the empirical analysis.

3. Empirical analysis

The empirical analysis commences by illustrating what one would obtain when applying the original F–H and Sachs' approaches to a dataset that covers more countries as well as a more recent sample period than that employed in many of these earlier studies. The starting point is the standard F–H cross-section regression model written as:

$$I_i = \alpha_0 + \alpha_1 S_i + u_i \tag{1}$$

where I is investment, S is domestic savings (both expressed as a percentage of GDP), u_i is the equation error term, and $i = 1, \ldots, N$ indicates the 25 countries used for estimation. The savings retention coefficient is denoted by α_1. Our estimates of Eq. (1) for each year from 1970 to 2011 are reported in columns 2 and 3 in Table 1. In addition to this, we also estimate the two model specifications advocated by Sachs (1981), namely the one in which the current account is regressed against an intercept and investment (columns 4 and 5 in Table 1), that is:

$$CA_i = \beta_0 + \beta_1 I_i + \varepsilon_i, \tag{2}$$

as well as that in which an intercept and gross savings are used as independent variables (columns 6 and 7 in Table 1), that is:

$$CA_i = \gamma_0 + \gamma_1 S_i + \xi_i. \tag{3}$$

While zero slope coefficients in Eqs. (2) and (3) are consistent with zero capital mobility, the respective slopes of −1 and 1 are consistent with perfect capital mobility. Focussing first on the estimates of the F–H specification, the salient feature that emerges from the analysis is that the estimate of the slope coefficient exhibits a declining trend passing from a highly statistically significant value of 0.564 in 1972 to a statistically insignificant value of −0.048 in 2006. Generally, the period 1995–2011 is characterised by insignificant savings retention coefficients suggesting that capital has in fact been perfectly mobile in the latter part of the study period. Given that most of the sample of countries are members of the European Union who had relaxed their remaining capital controls by the early 1990s, these results confirm a subsequent increase in capital mobility. The estimates of the slope coefficient that we obtained for the 1970s are not as high as the original estimate in excess of 0.826 reported by Feldstein (1983). The last row in Table 1 reports the results that are obtained when the investment and gross savings ratios are averaged over the whole sample period, which is the approach adopted in the seminal contribution by Feldstein and Horioka (1980) and later on by Feldstein (1983). The estimate of the slope coefficient that results from using the data averaged through time is 0.156, which is not statistically significant at the 5% level, and therefore supports the view of perfect international capital mobility. This finding differs from the earlier results reported by Sinn (1992), who suggests an upward bias in the savings retention coefficient derived from Feldstein–Horioka regressions involving long-term averages.

We now turn to the Sachs regression model as expressed by Eqs. (2) and (3), which can be derived from the above F–H equation. When the current account is regressed against investment, the estimate of the slope coefficient has the expected negative sign in all years, but it is rarely statistically significant. By contrast, in the current account regressions that include savings as the independent variable, the slope coefficient has the expected positive sign approaching the value of one as time progresses. The insignificance of investment points towards capital immobility and suggests that current account deficits (surpluses) are the result of reduced (increased) domestic savings and not investment booms (slumps).[3] The last row in Table 1 reports the results that are obtained when the current account, savings and investment ratios are averaged over the whole sample period. These results differ from Sachs (1981) who concludes that current account deficits are the result of investment booms based on a 1960–1979 study period, and Krol (1996) who finds both investment and savings ratios to be significant based on a 1962–1990 study period.

As indicated above, the previous year-by-year results are intended to serve only as an illustration, since coefficients may be biased for several reasons. First, year-by-year cross sectional regressions assume that the error term in the regression of one year is independent from that in the other years. However, many studies apply cointegration and error-correction techniques (see, inter alia, Miller, 1988; Gulley, 1992; Leachman, 1991; Sinha & Sinha, 2004) and show both short- and long-term effects between savings and investment. Second, macroeconomic aggregates across countries exhibit significant heterogeneity along both the

[2] Feldstein and Horioka (1980) exclude France, Iceland, Luxembourg, Mexico, Norway, Portugal, Spain, Switzerland and Turkey from their estimations. Furthermore, West Germany in the original F–H sample is replaced by Germany.

[3] In fact, further results (not reported here) indicate that for the years since the mid-1990s, the null hypothesis of a unity slope coefficient on savings cannot be rejected.

Table 1
Annual investment and current account regressions 1970–2011.

Year	Investment		Current account		Current account	
	S_i	R^2	I_i	R^2	S_i	R^2
1970	0.362‡	0.306	−0.154	0.014	0.638‡	0.579
1971	0.550‡	0.501	−0.090	0.010	0.450‡	0.401
1972	0.564‡	0.479	−0.152	0.029	0.436‡	0.353
1973	0.471‡	0.367	−0.222	0.045	0.529‡	0.422
1974	0.227	0.125	−0.447	0.086	0.773‡	0.625
1975	0.406†	0.246	−0.396	0.123	0.594‡	0.410
1976	0.360†	0.205	−0.431	0.129	0.640‡	0.449
1977	0.438†	0.214	−0.512†	0.231	0.562‡	0.309
1978	0.315	0.134	−0.576†	0.221	0.685‡	0.421
1979	0.221	0.070	−0.681‡	0.256	0.779‡	0.486
1980	0.322†	0.208	−0.356	0.074	0.678‡	0.537
1981	0.239	0.098	−0.589†	0.183	0.761‡	0.525
1982	0.281	0.125	−0.556†	0.183	0.719‡	0.483
1983	0.171	0.049	−0.713†	0.241	0.829‡	0.549
1984	0.242	0.133	−0.450	0.093	0.758‡	0.601
1985	0.242	0.113	−0.535	0.144	0.758‡	0.554
1986	0.412‡	0.313	−0.242	0.044	0.588‡	0.480
1987	0.429‡	0.354	−0.175	0.024	0.571‡	0.493
1988	0.474‡	0.393	−0.170	0.026	0.526‡	0.444
1989	0.361†	0.249	−0.311	0.063	0.639‡	0.510
1990	0.293†	0.196	−0.331	0.056	0.707‡	0.587
1991	0.334‡	0.260	−0.223	0.028	0.666‡	0.582
1992	0.238†	0.160	−0.328	0.043	0.762‡	0.661
1993	0.243†	0.174	−0.286	0.033	0.757‡	0.670
1994	0.201	0.119	−0.409	0.061	0.799‡	0.681
1995	0.129	0.055	−0.574	0.095	0.871‡	0.727
1996	0.116	0.053	−0.540	0.072	0.884‡	0.767
1997	0.126	0.062	−0.508	0.066	0.874‡	0.761
1998	0.094	0.039	−0.579	0.072	0.906‡	0.794
1999	0.099	0.061	−0.379	0.024	0.901‡	0.844
2000	−0.033	0.009	−1.279†	0.162	1.033‡	0.902
2001	0.034	0.008	−0.776	0.084	0.966‡	0.862
2002	0.034	0.007	−0.795	0.095	0.966‡	0.850
2003	0.066	0.027	−0.587	0.054	0.934‡	0.848
2004	0.030	0.006	−0.804	0.090	0.970‡	0.862
2005	0.015	0.001	−0.910	0.121	0.985‡	0.854
2006	−0.048	0.011	−1.228†	0.242	1.048‡	0.841
2007	−0.025	0.005	−1.189†	0.159	1.025‡	0.888
2008	0.027	0.009	−0.679	0.037	0.973‡	0.920
2009	0.055	0.026	−0.531	0.033	0.945‡	0.887
2010	0.016	0.002	−0.867	0.082	0.984‡	0.892
2011	0.032	0.008	−0.759	0.072	0.968‡	0.876
1970–2011	0.156	0.103	−0.344	0.031	0.844‡	0.769

Notes: All regressions, performed in Stata 11, include an intercept term (not reported to save space). Estimates in the last row are calculated using the investment and gross savings ratios averaged over the whole sample period. † and ‡ denote statistical significance at the 5% and 1% levels, respectively.

cross-section and time dimensions. Third, the static nature of the analysis makes it difficult to determine whether domestic savings truly end up as domestic investment, or not.

Thus, in an attempt to exploit the informative power of the variables that have been collected across countries and over time, we consider the information set as a panel of data. Such a modelling approach might turn out to be useful to shed some light on the identification and measurement of effects that cannot be detected using either cross-section or time-series data only. Once the data panel has been assembled, we follow a modelling approach based on the recent tools from the econometric literature on heterogeneous panels. The underlying idea in this strand of the literature is to exploit the fact that the dataset consists of a not so large number of countries ($N = 25$) observed over a much longer time period ($T = 42$). Having a so-called large T panel opens up the possibility of estimating separate regressions for each country, and then the averages of the resulting coefficients can be calculated over these countries (along with their corresponding standard errors).

Our empirical analysis starts off by examining the time series properties of the macroeconomic aggregates under consideration. For this, we first examine whether innovations (shocks) to the current account, investment and savings ratios are cross-sectionally independent. To do this, we estimate the Pesaran (2004) general diagnostic test for cross section dependence in panels, denoted as the CD statistic. Monte Carlo simulation results reported by this author indicate that the CD test performs well in terms of its size and power properties in small samples, being also robust to the presence of structural breaks and unit roots. To implement the CD test, ADF regressions are fitted to each cross section unit i separately, using p lags of the dependent variable, and the resulting residuals of the individual series are denoted as $\hat{e}_{it}$. The motivation behind this initial stage is to capture any serial correlation in the individual time series i. Then, the cross-correlation coefficient between the residuals of cross section units i and j is computed as:

$$\hat{\rho}_{ij} = \frac{\sum_{t=1}^{T} \hat{e}_{it}\hat{e}_{jt}}{\left(\sum_{t=1}^{T} \hat{e}_{it}^2\right)^{1/2}\left(\sum_{t=1}^{T} \hat{e}_{jt}^2\right)^{1/2}}. \tag{4}$$

Finally, the CD statistic is calculated as:

$$CD = \sqrt{\frac{2T}{N(N-1)}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}\right) \sim N(0,1). \tag{5}$$

Table 2 summarises the results of applying the CD statistic to the 25 current account, investment and savings ratios using up to $p = 6$ lags in the ADF(p) regressions, when both country-specific intercepts and trends are included in the test regressions.[4] As can be seen from the table, the null hypothesis that the series innovations are cross-sectionally independent is strongly rejected. This finding, which is irrespective of whether one includes country-specific intercepts only and/or shorter lag lengths, provides a motivation for employing panel data techniques as a tool of analysis. Moreover, these initial findings are consistent with innovations (shocks) to the current account, investment and savings ratios being cross-sectionally dependent which underlie the appropriateness of conducting our analysis within a panel data framework.

Next, we study the order of integration of the macroeconomic aggregates under consideration. For this, one of the most commonly used testing procedures available in the literature is the Pesaran (2007) CIPS test, which tests the joint null hypothesis of a unit root against the alternative of at least one stationary series in the panel. Although the CIPS test successfully overcomes the limitations of the Im, Pesaran, and Shin (2003) (IPS) test in the presence of cross-sectional dependence, care must be exercised when interpreting the results. Indeed, due to the heterogeneous nature of the alternative hypothesis in both the IPS and CIPS tests, the rejection of the null hypothesis may occur when only a fraction of the series in the panel is stationary. Furthermore, neither the IPS nor the CIPS test allows for the possibility of structural breaks which, if erroneously omitted, can cause deception in time series testing, and whose effects do not disappear simply because one is using panel data. Lack of a careful investigation of potential structural breaks may lead to misspecification.

Taking the abovementioned aspects into consideration, we follow Hadri (2000) and Hadri and Rao (2008) who develop a panel version of the univariate Kwiatkowski, Phillips, Schmidt, and Shin (1992) stationarity test, so that under the null hypothesis all series are stationary, while under the alternative some (but not necessarily all) of the series have a unit root. Notice that formulating the null hypothesis in this fashion is advantageous because if the null is not rejected, then one may conclude that all the series in the panel are stationary. Furthermore, the testing approach permits us to accommodate one-time endogenously determined structural breaks, serial correlation and cross-section dependence (the latter two through the use of an autoregressive bootstrap procedure).

Hadri (2000) studies the following models:

$$y_{it} = \alpha_i + r_{it} + \varepsilon_{it}, \tag{6}$$

$$y_{it} = \alpha_i + r_{it} + \beta_i t + \varepsilon_{it}. \tag{7}$$

where y_{it} denotes the observed values of the variable of interest, namely the current account, investment or savings for country i at time t, $i = 1, \ldots, N$, $t = 1, \ldots, T$, r_{it} is a random walk, $r_{it} = r_{it-1} + u_{it}$, and ε_{it} and u_{it} are mutually independent normal distributions. Also, ε_{it} and u_{it} are independent and identically distributed (i.i.d.) across i and over t, with $E[\varepsilon_{it}] = 0$, $E[\varepsilon_{it}^2] = \sigma_{\varepsilon,i}^2 > 0$, $E[u_{it}] = 0$ and $E[u_{it}^2] = \sigma_{u,i}^2 \geq 0$. Within this framework, the null hypothesis that all the individual series in the panel are stationary is $H_0: \sigma_{u,i}^2 = 0$, where $i = 1, \ldots, N$. The alternative hypothesis that some (but not all) of the individual series have a unit root is $H_1: \sigma_{u,i}^2 > 0$, where $i = 1, \ldots, N_1$; and $\sigma_{u,i}^2 = 0$, where $i = N_1 + 1, \ldots, N$.

The models in Eqs. (6) and (7) are used to test for the level and trend stationarity, respectively. In the case of savings, for example, the inclusion of a time trend might be justified by the presence of demographic factors that can affect aggregate consumption and savings behaviour. In a subsequent paper, Hadri and Rao (2008) further extend this setup by considering models that accommodate a (one-time) structural break under the null hypothesis:

$$y_{it} = \alpha_i + r_{it} + \delta_i D_{it} + \varepsilon_{it}, \tag{8}$$

$$y_{it} = \alpha_i + r_{it} + \delta_i D_{it} + \beta_i t + \varepsilon_{it}, \tag{9}$$

$$y_{it} = \alpha_i + r_{it} + \beta_i t + \gamma_i DT_{it} + \varepsilon_{it}, \tag{10}$$

$$y_{it} = \alpha_i + r_{it} + \delta_i D_{it} + \beta_i t + \gamma_i DT_{it} + \varepsilon_{it}, \tag{11}$$

[4] The results of the CD test are based on computer code developed by the authors using the Rats 8.2 computer econometric software.

Table 2
CD cross-section independence test.

Panel	CD test	*p*-value
Current account	6.322	[0.000]
Investment	18.386	[0.000]
Savings	21.745	[0.000]

Notes: The CD test is calculated including a constant and a trend. The number of lags of the dependent variable included in the ADF-type regressions is 6. The *p*-values are based on the standard normal distribution.

where, in addition to the terms already defined, D_{it} and DT_{it} are dummy variables to specify the type of structural break, which are in turn defined as:

$$D_{it} = \begin{cases} 1, & \text{if} \quad t > T_{B,i} \\ 0 & \text{otherwise} \end{cases} \tag{12}$$

and

$$DT_{it} = \begin{cases} t - T_{B,i}, & \text{if} \quad t > T_{B,i} \\ 0, & \text{otherwise} \end{cases}, \tag{13}$$

where $T_{B,i}$ denotes the time of occurrence of the structural break for individual i. Also, $T_{B,i} = \omega_i T$, where $\omega_i \in (0, 1)$ indicates the fraction of the break point relative to the whole sample period for individual i. The parameters δ_i and γ_i measure the extent (or magnitude) of the structural break, and allow for the possibility of different breaking dates across the individuals in the panel.

The models in Eqs. (8) to (11) characterise the different effects that a structural break can have on the deterministic component of a series. More specifically, in Eq. (8) there is an intercept term and the shift occurs in the level of the series. In Eq. (9) there are intercept and linear trend terms, and the shift affects the former (but not in the latter). In Eq. (10) there are intercept and linear trend terms, and the change occurs in the latter (but not in the former). Lastly, Eq. (11) incorporates intercept and linear trend terms, and allows for a change in both the level and the slope of the series.

Hadri and Rao (2008) apply a systematic procedure to find the model that best characterises each series. Specifically, they begin by determining the date of the break point endogenously, which involves estimating for each cross section unit in the panel and for each model the break date, $\hat{T}_{B,i,k}$. This can be accomplished by minimising, with respect to $0 < \omega_i < 1$, the residual sum of squares (RSS) from the relevant model under the null hypothesis, where $i = 1, \ldots, N$ denotes the series, and $k = 1, 2, \ldots, 6$ refers to the models stated in Eqs. (6) to (11). Then, given $\hat{T}_{B,i,k}$, for each series, i, the preferred model, k, is chosen by minimising the Schwarz information criterion.[5]

Letting $\hat{\varepsilon}_{it}$ be the residuals that result from estimating the chosen model (with or without a break), then the idea is to use $\hat{\varepsilon}_{it}$ to calculate the KPSS statistic, for which a critical issue is how to estimate the so-called long-run variance. Here, we follow Carrion-i-Silvestre and Sansó (2006) who, based on Monte Carlo simulation results, recommend estimation of this key parameter applying the boundary condition rule advocated by Sul, Phillips, and Choi (2005). As highlighted earlier, an important issue here is related to the fact that, similar to other first generation panel unit root tests (such as IPS), following the terminology in Breitung and Pesaran (2008), the Hadri (2000) and Hadri and Rao (2008) tests suffer from severe size distortions in the presence of cross-sectional dependence, the magnitude of which increases as the strength of the cross-sectional dependence increases; see Giulietti, Otero, and Smith (2009). Thus, an important feature of our analysis is that we allow for the presence of both serial correlation and cross section dependence by means of the implementation of an autoregressive-based bootstrap.[6]

Table 3 presents the results of the KPSS tests for stationarity around a trend. Generally speaking, the KPSS results offer mixed support for the view that the current account, investment and savings series can be characterised as stationary processes over the study period. In fact, in some cases the results contradict the rules of the linear combinations of integrated series (see Granger, 1981). Thus, for instance, for Australia, Austria, Canada, Luxembourg and Netherlands, both investment and savings are I(0), while the current account, which is simply the difference between the two, is I(1). A similar contradictory finding occurs for Denmark, New Zealand, Sweden and Switzerland, where savings is I(1), investment is I(0), and the current account is not I(1) but I(0); see also Portugal, where investment is I(1), while savings and the current account are both I(0). However, and in what can be thought of as an interesting twist of results, a rather different picture emerges when these macroeconomic aggregates are analysed within a panel context, and allowance is made for cross sectional dependence (which, as documented in Table 2, is a feature that characterises the data used in this paper). Indeed, the results reported in the first row of Table 4 indicate that savings, investment and the current account can all be best described as trend-stationary processes. If one instead tests for mean stationarity, results not reported here also indicate failure to reject the stationarity null at the 5% (but not 10%) significance level.

[5] Notice that when implementing the test the models in Eqs. (6) and (7) are estimated only once, since they do not include the dummy variables D_{it} and DT_{it}.
[6] See Hadri and Rao (2008) for a description of how to perform the autoregressive-based bootstrap; the results reported here are based on our own computer code developed in Rats 8.2.

Table 3
KPSS tests for trend stationarity.

Country	Current account		Investment		Savings	
	Lag	Statistic	Lag	Statistic	Lag	Statistic
Australia	6	0.209†	1	0.058	1	0.077
Austria	6	0.167†	1	0.043	1	0.082
Belgium	6	0.225‡	2	0.085	5	0.290‡
Canada	6	0.204†	2	0.085	1	0.041
Denmark	1	0.037	2	0.114	6	0.242‡
Finland	6	0.099	2	0.096	2	0.084
France	1	0.034	2	0.111	1	0.059
Germany	1	0.109	2	0.125	1	0.093
Greece	3	0.125	2	0.106	1	0.102
Iceland	2	0.080	5	0.045	1	0.065
Ireland	6	0.216‡	6	0.140	6	0.173†
Italy	2	0.081	2	0.111	1	0.036
Japan	1	0.043	6	0.115	1	0.028
Luxembourg	1	0.167†	1	0.052	1	0.124
Mexico	1	0.065	2	0.099	1	0.030
Netherlands	4	0.344‡	2	0.117	1	0.045
New Zealand	1	0.070	2	0.077	6	0.210†
Norway	2	0.109	2	0.079	1	0.091
Portugal	5	0.080	6	0.180†	2	0.115
Spain	3	0.136	2	0.087	2	0.089
Sweden	1	0.050	2	0.078	6	0.418‡
Switzerland	1	0.061	2	0.101	5	0.382‡
Turkey	1	0.062	1	0.072	1	0.060
United Kingdom	5	0.172†	6	0.283‡	2	0.116
United States	6	0.277‡	6	0.180†	6	0.149

Notes: Lag indicates the optimal number of lags used to account for residual serial correlation when computing the KPSS statistic, as determined by a General-To-Specific (GETS) algorithm with $p = 6$ lags. GETS involves fitting an AR(p) model to the de-trended (de-meaned) series, and testing if the last estimated coefficient is statistically different from zero at some level of significance, let us say 5%; if the coefficient is not significant then the order of autoregression is reduced by one, until the last estimated coefficient is found to be significant. Qualitatively similar results are obtained when $p < 6$. † and ‡ denote statistical significance at the 5% and 1% levels, respectively, based on finite sample critical values calculated from the response surfaces in Sephton (1995). The long-run variance required to calculate the KPSS statistic is consistently estimated using the new boundary condition rule put forward by Sul et al. (2005).

The second row in Table 4 shows that the finding of panel stationarity is also obtained when performing the Hadri and Rao (2008) test. As pointed out earlier, this testing procedure allows us to account for the presence of one-time structural breaks which, as can be seen in Table 5, appears to have occurred at some point in time in the savings and investment series under consideration, with the exception of the series of savings in Luxembourg.[7] Overall, the occurrence of the breaks seems to be specific to each country and does not necessarily affect the savings and investment series within a country in a similar fashion. The fact that the date of the break is being estimated, rather than exogenously imposed, is in sharp contrast with the modelling approach adopted by Rao et al. (2010), where the influence of both the Bretton Woods and Maastricht agreements on 13 OECD countries is examined. In the case of our sample, the majority of OECD countries experience a break in at least one of the savings and investment series prior to 1990.

The stationarity finding is in agreement with economic intuition, as all the series are being measured as a percentage of GDP and are therefore bounded in the [0, 1] interval.[8] Equally important, the stationarity result precludes the need to test for the possibility of cointegration among investment and savings. Gundlach and Sinn (1992) have previously argued that a mean stationary current account points to a stable long-run cointegrating relationship between non-stationary savings and investment and so no linkage between international capital markets. Given the stationarity of savings and investment found here, the Gundlach and Sinn approach has much less relevance to our assessment of capital mobility.

Available estimators to investigate capital mobility in a heterogeneous panel setup include the mean group (MG) estimator studied in Pesaran and Smith (1995), and the more recent cross correlated effects mean group (CCEMG) estimator put forward in Pesaran (2006). Here, we focus on the latter as it allows for the presence of cross-sectional dependence and endogeneity. As is well known in the literature, the relationship between domestic savings and investment may be driven by the factors such as productivity shocks, demographic changes, and government policies, among others, which have nothing to do with capital mobility. Similarly, in the presence of large offsetting capital inflows and outflows, the slope coefficient in the investment–savings

[7] Visual inspection of the current account series that are calculated with the detrended versions of the savings and investment series suggests the absence of noticeable structural breaks.

[8] From an economic point of view, Cerrato et al. (2013) argue that a justification for the average propensity to consume being trended/trend stationary can be in terms of an income distribution that changes over time. From an econometric perspective, testing for trend stationarity (rather than for mean stationarity) can be justified on the ground that this specification provides a better fit of the deterministic component of the variables under examination; see also Ho (2002). On this point, it is perhaps useful to recall that the key difference between a stochastic trend and a deterministic trend is that in the former the series is not mean reverting, whereas in the latter it fluctuates around a non-constant mean.

Table 4
Panel stationarity tests.

Test	Current account		Investment		Savings	
	Statistic	Bootstrap p-value	Statistic	Bootstrap p-value	Statistic	Bootstrap p-value
Hadri	7.340	[0.921]	4.436	[0.986]	7.242	[0.439]
Hadri–Rao	0.591	[0.999]	6.287	[0.898]	10.106	[0.201]

Notes: Hadri is the panel stationarity test including constant and trend but no structural breaks. Hadri and Rao is the panel stationarity test allowing for the presence of one-time endogenously determined structural breaks. For both tests the p–values are based on 2000 replications of an autoregressive bootstrap. The bootstrap procedure allows us to account for the presence of serial correlation and cross-section dependence.

equation could be one regardless of the degree of capital mobility; see the discussion in Georgopoulos and Hejazi (2009). A similar difficulty arises when one tries to infer the degree of capital mobility based on the relationship between the current account and the savings and investment ratios. In the presence of large offsetting capital inflows and outflows, current account and savings (investment) may appear to be uncorrelated even there is a perfect capital mobility. The structure of the model CCEMG estimator proposed by Pesaran (2006) is general enough to allow for the possibility that an unobserved common factor could be correlated with the individual (country) specific regressors. Consequently, although we do not formally attempt to address the endogeneity issue described above with regard to specific variables, the multifactor error structure underpinning the CCEMG estimator permits us to accommodate the potential endogeneity that arises when the unobserved common factors affect both the dependent and independent variables.

The CCEMG estimator is based on the equation:

$$I_{it} = \alpha_{0i} + \alpha_{1i} S_{it} + \alpha_{2i} \bar{I}_t + \alpha_{3i} \bar{S}_t + e_{it}, \tag{14}$$

where $\bar{I}_t = {}^1/_N \sum_i I_{it}$ and $\bar{S}_t = {}^1/_N \sum_i S_{it}$ denote the cross section means of investment and domestic savings (as a percentage of GDP), respectively. The idea behind Eq. (14) is that augmenting the regression with the cross-sectional means of the dependent and independent variables helps to capture the unobserved common factor (such as global shocks) which induces the cross-sectional dependence. It ought to be noticed that the effect of the unobserved common factor, as captured by the terms $\bar{I}_t$ and $\bar{S}_t$, is permitted to differ across countries. The CCEMG estimator of the savings retention coefficient is thus given by the simple average of the corresponding estimated coefficients in Eq. (14), that is $\hat{\alpha}_{1,CCEMG} = {}^1/_N \sum_i \hat{\alpha}_{1i}$. Of course, for the Sachs current account regressions, Eq. (14) must be modified by changing the dependent and independent variables accordingly.

Table 5
Estimated models and structural breaks.

Country	Investment		Savings	
	Model	Break date	Model	Break date
Australia	11	1990	10	1999
Austria	10	1998	11	1988
Belgium	8	1992	11	1977
Canada	8	1982	10	1993
Denmark	11	2005	8	1982
Finland	11	2005	11	1997
France	11	1993	11	1995
Germany	9	1988	9	1986
Greece	11	1987	10	1981
Iceland	11	1983	11	1986
Ireland	11	1985	11	1981
Italy	11	1988	10	1983
Japan	11	1978	11	1986
Luxembourg	8	1988	6	No
Mexico	11	1984	11	1978
Netherlands	11	1993	11	1977
New Zealand	9	1999	11	1986
Norway	8	2009	11	1996
Portugal	11	1981	11	1987
Spain	11	1992	9	1999
Sweden	11	1980	9	1975
Switzerland	9	1989	11	1985
Turkey	8	1980	11	1984
United Kingdom	9	1999	11	1984
United States	11	1992	11	1998

Notes: The columns labelled Model indicate the chosen model specifications, as postulated in Eqs. (6) to (11).

In a related paper, Bebczuk and Schmidt-Hebbel (2010) apply the Pesaran, Shin, and Smith (1999) pooled MG (denoted PMG) estimator for the F–H savings retention coefficient. This approach may be viewed as a restricted version of the CCEMG estimator outlined earlier since it does not include the cross-sectional means of the dependent and independent variables. Roughly speaking, the PMG estimator is derived after estimating individual autoregressive distributive lag (ARDL) models for each individual in the panel, imposing homogeneity in the long-run coefficients, but allowing the short-run coefficients and error variances to differ across individuals. A crucial issue here is that while Bebczuk and Schmidt-Hebbel (2010) recognise the importance of cross sectional dependence, they do not formally test for it. Instead, they employ data that has been previously demeaned (by subtracting annual cross-country averages) in an attempt to eliminate common international systematic factors. However, Maddala and Wu (1999) point out that demeaning the data does not always succeed in eliminating those factors that are common to all individuals.[9]

Table 6 summarises the results of the CCEMG estimator as applied to two versions of the dataset, namely one with the original data series (i.e. without structural break) and the other one with the estimated broken-trend deterministic component removed from them (i.e. with structural break). We begin by briefly referring to the country-level results which confirm considerable heterogeneity across countries in terms of the savings–investment association, regardless of whether the estimations are performed with or without structural breaks; the aspect of the heterogeneous coefficients has been highlighted by Caporale, Panopoulou, and Pittis (2005) and others.

Putting aside for the moment the presence of structural breaks, Table 6 shows that the savings retention coefficient is somewhat small compared to F–H, being equal to 0.334, though statistically different from zero so that perfect capital mobility is ruled out.[10] This finding can be compared with Coakley et al. (2004), who examine a smaller sample of countries over the much shorter study period covering 1980Q1–2000Q4. According to these authors, the MG and cross-sectionally augmented MG estimates are characterised by small coefficients such they are unable to reject the null of a zero slope implying perfect capital mobility. As for Bebczuk and Schmidt-Hebbel (2010), they find a long-term PMG estimator between savings and investment equal to 0.75. Nonetheless, following the cautionary point raised by Maddala and Wu (1999), it remains to be seen how well demeaning their data has addressed the cross section dependence that characterises the macroeconomic aggregates being studied.

Turning to the Sachs current account models when the original data are employed, which are not considered by either Coakley et al. (2004) or Bebczuk and Schmidt-Hebbel (2010), the estimates of the two slope coefficients in the two current account equations are of similar magnitude and with opposite signs.[11] In short, the heterogeneous panel data results reported in Table 6 highlight the importance of allowing for heterogeneity (along the cross-section and time dimensions) and cross section dependence when modelling the association among the variables under consideration. These latter results provide stronger evidence of capital mobility than the year-by-year results reported in Table 1; indeed, it is noticeable that in the equation of the current account against investment (savings) the estimate of the slope coefficient of −0.676 (0.666) is statistically different from zero, being closer, though not statistically equal, to the value of −1 (1) that is consistent with perfect capital mobility.

Let us now turn our attention to the results when the estimated broken-trend component is removed from the data series prior to the implementation of the CCEMG estimator. In the case of the F–H regression the point estimate of the slope coefficient is now much smaller (i.e. 0.179). This is consistent with the findings obtained by Rao et al. (2010) in the sense that the Bretton Woods and Maastricht seem to have weakened the F–H puzzle in their study. This is an interesting finding because in our analysis, the structural breaks are not imposed from the outset but endogenously determined by the data series themselves. Once we allow for structural breaks, the savings retention coefficient is lower than previously found both in our study here and in most of the earlier studies. Given that we find capital mobility to be very high though not perfect, an implication from this is the likely ease by which a financial crises can spread between countries. The recent experiences surrounding the global financial crisis come to mind here, so in this respect our findings are in line with expectations. The presence of home-bias in the allocation of domestic savings is most probably reflected in a savings retention coefficient that is significantly different from zero as agents maintain some degree of wariness of investing in foreign countries.

As to the Sachs' current account model specifications, the estimated slope coefficients are (as expected) of similar magnitude and with the different sign. Interestingly, once the (one-time) structural breaks are removed from the underlying series, the resulting point estimate of the slope coefficient is equal to −0.827 (0.821) when the current account is regressed on investment (savings). The fact that the estimated coefficients are respectively closer to −1 and 1 is consistent with our earlier finding for the FH specification, in the sense that once the structural breaks are accommodated in the analysis, there appears to be more evidence of improved capital mobility. Our findings are more consistent with Krol (1996), who finds both the investment and savings ratios to be significant in driving the current account, than with Sachs (1981), who concludes that current account deficits are the result of investment booms.

In light of these findings of high capital mobility, it is important to reflect on the policy implications that can arise. The well-known Trilemma of international finance argues that under perfect capital mobility, a given country cannot have an

[9] Another difference between our work and that of Bebczuk and Schmidt-Hebbel (2010) is that they do not examine the time series properties of the macroeconomic variables, nor whether they are cointegrated. Therefore, it is not clear whether the error correction models that they estimate are balanced, in the sense that the equations contain variables with the same order of integration on both sides.

[10] Although the results from CCEMG estimation include country-specific time trends in order to account for home bias in the allocation of domestic savings, as recommended by Georgopoulos and Hejazi (2005), they are qualitatively the same if these country-specific terms are omitted from the regressions.

[11] It is perhaps worth mentioning that the results of the Sachs' approach in Table 4 are qualitatively similar if the country-specific trend terms are omitted from the regressions.

Table 6
CCEMG heterogeneous panel estimators.

Country	I_t on S_t		CA_t on I_t		CA_t on S_t	
	Without break	With break	Without break	With break	Without break	With break
CCEMG estimator	0.334‡	0.179‡	−0.676‡	−0.827‡	0.666‡	0.821‡
Australia	0.601‡	0.428‡	−0.163	−0.448‡	0.399‡	0.572‡
Austria	0.369†	−0.116	−0.702‡	−1.112‡	0.631‡	1.116‡
Belgium	0.680‡	0.176	−0.095	−0.877‡	0.320‡	0.824‡
Canada	0.048	0.346	−0.891‡	−0.793‡	0.952‡	0.654‡
Denmark	0.050	0.210	−0.930‡	−0.753‡	0.950‡	0.790‡
Finland	−0.008	−0.087	−1.007‡	−1.302‡	1.008‡	1.087‡
France	0.380†	0.009	−0.653‡	−0.962‡	0.620‡	0.991‡
Germany	0.601‡	−0.165	−0.379‡	−1.135‡	0.399‡	1.165‡
Greece	0.353†	0.134	−0.693‡	−0.915‡	0.647‡	0.866‡
Iceland	−0.113	0.095	−1.108‡	−0.942‡	1.113‡	0.905‡
Ireland	0.444‡	0.256	−0.540‡	−0.881‡	0.556‡	0.744‡
Italy	−0.250	0.290‡	−1.362‡	−0.439	1.250‡	0.710‡
Japan	0.732‡	0.134	−0.104	−0.813‡	0.268‡	0.866‡
Luxembourg	−0.037	0.252	−1.151‡	−0.705‡	1.037‡	0.748‡
Mexico	−0.044	0.671‡	−1.201‡	−0.527‡	1.044‡	0.329
Netherlands	0.313	0.016	−0.799‡	−0.980‡	0.687‡	0.984‡
New Zealand	0.514†	0.448‡	−0.806‡	−0.592‡	0.486	0.552‡
Norway	−0.011	0.214†	−1.013‡	−0.461	1.011‡	0.786‡
Portugal	0.398‡	0.496‡	−0.434†	−0.571‡	0.602‡	0.504‡
Spain	0.721‡	0.206	−0.662‡	−0.618‡	0.279	0.794‡
Sweden	0.616‡	−0.118	−0.622‡	−1.145‡	0.384†	1.118‡
Switzerland	0.632‡	0.360	−0.544‡	−0.843‡	0.368†	0.640‡
Turkey	0.501‡	0.099	0.091	−0.975‡	0.499‡	0.901‡
United Kingdom	0.127	0.041	−0.839‡	−0.977‡	0.873‡	0.959‡
United States	0.730‡	0.087	−0.300‡	−0.911‡	0.270†	0.913‡

Notes: When there is no break, the CCEMG estimator is computed including intercept, trend and the cross sectional means of the dependent and independent variables in the test regressions. When there is break the test regressions include the cross sectional means of the dependent and independent variables, but no deterministic component (as it has been previously removed from each series). The estimated coefficients on these additional variables are not reported to save space. These estimations were performed with Stata 11. † and ‡ denote statistical significance at the 5% and 1% levels, respectively.

independent monetary policy with a fixed exchange rate, or that the exchange rate must float if that country is to have an independent monetary policy. Given that the Euro is a floating currency, the Euro members from our sample are in a position to have a (common) independent monetary policy with respect to non-Euro countries. In the case of fiscal policy, there are policy implications in terms of how the budget and current account deficits might be related. For example, the Mundell–Fleming model suggests that a fiscal expansion could worsen the current account deficit through an increase in domestic interest rates, capital inflows and an appreciation of the exchange rate. Policymakers may need to keep a close watch on the impact on the external balance. If capital is perfectly mobile, this will encourage policymakers to more closely examine the microeconomic considerations and incentives behind the movement of financial resources across borders. Further to this, the Sachs' approach points to the role of increased domestic savings as a means of improving the external balance, so policy analysis might explore the microeconomic incentives behind improving the rate of domestic savings used as a means for funding domestic investment.

4. Concluding remarks

This paper re-examines two complementary views of international capital mobility, namely the one of Feldstein and Horioka (1980), according to which the increased movement of capital flows should be reflected in low correlations between domestic savings and investment, and the one of Sachs (1981), which instead focuses on the relationships between the current account and savings and investment ratios. A key message from our paper is that capital mobility for the OECD countries is greater than hitherto thought. Indeed, using a sample of 25 OECD countries, and employing the original regression specification used by Feldstein–Horioka estimated over a (more recent) time period, we find that in contrast to earlier work the use of time-averaged data does not play down the extent of capital mobility. Further analysis using Sachs' approach to test for capital mobility supports this finding, and also suggests a lack of capital mobility reflected in only domestic savings and not in investment driving current account balances. We address the issue of potential cross-sectional and time aggregation bias through an analysis of the data employing estimation techniques from the literature on heterogeneous panels. Such an approach enables us to exploit the information power contained in both the cross-section and time-series dimensions of the data, also allowing us to adequately address cross-sectional dependence and potential endogeneity. Not only do these results support the view of stronger international capital mobility, but they also highlight the role of both savings and investment as factors that helps explain the behaviour of the current account.

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