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A Pairwise-Based Approach to Examine the Feldstein-Horioka Condition of International Capital Mobility

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Abstract

We propose a pairwise procedure to test the Feldstein-Horioka condition of capital mobility. In contrast to the existing approach, we explicitly examine the relationship between domestic investment and foreign savings rather than domestic savings. In terms of addressing the Feldstein-Horioka puzzle, our results based on a panel of OECD and emerging market economies initially suggests that the depth and extent of capital mobility remains generally limited, and that mobility has increased over the past twenty years. However, in contrast to existing studies, we find that capital mobility between Euro and EU pairs is more extensive than between pairs that involve other countries. If our sample is expanded to include emerging markets, we find that capital mobility has also increased though is weaker than for OECD economies. We provide additional insight in terms of consistency between our assessment of capital mobility based on the Feldstein-Horioka condition (a quantity approach) and a price approach based on real interest rate differentials.

Key words

Feldstein-Horioka
capital mobility
pairwise real interest rates.

JEL Classification

F3; F4; F6

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1 Introduction

In a paper that has stimulated a great deal of subsequent work in open economy macroeconomics, Feldstein and Horioka (1980) (FH) argue that in a world with perfect capital mobility, domestic savings will search for the capital markets with the highest real returns, with no effect whatsoever of domestic savings on domestic investment; see, inter alia, Frankel (1992) and Taylor (1996). Using the degree of correlation between domestic savings and investment in order to judge the extent of international capital mobility has proved to be controversial though. Consequently, there exists a large literature that explores the application of the FH test for capital mobility; see, inter alia, survey articles by authors such as Lapp (1996), Coakley et al. (1998) and Apergis and Tsoumas (2009) and references therein. A significant part of this literature dwells on the FH puzzle in that despite the relaxation of capital controls, capital mobility does not appear to be particularly high as reflected by the positive correlations between domestic savings and domestic investment reported in previous studies.

In this paper, we concern ourselves with two issues related to the insights provided by the FH test. Indeed, we see these issues very much as shortcomings with existing studies in the way that the FH hypothesis has been tested and results interpreted. First, the existing tests of capital mobility based on the FH hypothesis focus on the relationship between *domestic* saving and domestic investment. There is no explicit assessment of the relationship between *foreign* saving and domestic investment. Based on the degree of correlation between domestic investment and domestic saving or size of the savings retention coefficient, the existing approach is able to reflect on the depth of capital mobility. However, there is no insight into the extent of capital mobility in terms of the numbers of cases where foreign savings are significantly correlated with domestic investment. In order to address this issue, we employ a pairwise approach based on analysing a panel of all foreign saving and domestic investment pairs for a sample of OECD economies. This approach enables us to reflect on the strength of capital mobility, not just in terms of the size of the saving retention coefficient, but also on the proportion of cases for which domestic investment is more strongly correlated with foreign rather than domestic savings. Further to this, we assess the changing nature of capital mobility over time through estimation based on a moving window approach.

There exist bilateral tests of capital market integration which tend to be price-based approaches such as testing for covered interest parity. The novelty of this study is that we develop the FH quantity-based approach towards a framework that tests the bilateral relationships between domestic investment and foreign saving across countries. This pairwise approach, built on the theory that underpins the basic FH hypothesis, enables us to address the puzzling result that despite increased capital mobility, Feldstein and Horioka (1980) and others find a high correlation between domestic saving and investment. In fact, not only does the suggested pairwise approach enable us to estimate the degree to which foreign as opposed to domestic savings are correlated with domestic investment but, by focusing on the relationship between domestic investment and foreign saving, we are also in a position to offer additional new insight based on the extent of the saving-investment correlations across a large pairwise combination of countries. Our results based on a panel of OECD economies suggest that the extent of capital mobility has increased significantly over the past twenty years. However, in contrast to existing studies, we find that capital mobility between Euro and EU pairs is more extensive than between pairs that involve other countries. If our sample is expanded to include emerging markets, we find that capital mobility has also increased though is weaker than for OECD economies. This is in contrast to the noted puzzle that saving-investment correlations are lower for emerging markets as compared with advanced economies.

In a second important direction, we show how our pairwise approach enables us to benchmark FH capital mobility measurement against alternative capital mobility and financial integration measures such as the real interest rate differential. If the FH approach is consistent with such alternative measures, then those pairs of countries with less capital mobility between them, according to this approach, should also be those pairs characterised by a larger real interest rate differential. We find that this is indeed the case thereby offering support to the consistency of the FH approach as a method of measuring capital mobility.

The paper is organised as follows. Section 2 describes the empirical strategy. Section 3 describes the data. Section 4 summarises the time-series properties of the data. Section 5 reports the results from applying our proposed pair-wise test to examine the FH condition of capital mobility. Section 6 relates the results from the pair-wise FH condition of capital mobility with the observed differential between real interest rates. Section 7 concludes.

2 Empirical modelling strategy

In contrast to the existing literature that tests the FH hypothesis, our approach offers a more explicit assessment of the relationship between domestic investment and foreign saving. Following Lemmen and Eijffinger (1995) and others, the FH approach can be viewed as a quantitative approach to examining financial integration and capital mobility. Rather than focus explicitly on interest parity conditions [Frankel and MacArthur (1988) and Frankel (1989)], the FH approach pays attention to macroeconomic aggregates. In doing so, the FH approach has linkages with Balance of Payments theories. In a world where foreign capital moves freely, domestic investment can be funded by foreign saving via a current account deficit (equal to the gap between investment and domestic saving). Indeed, those countries who are "net savers" can lend to those who are "net borrowers" where inter-temporal borrowing and lending achieves a faster accumulation of investment and a more efficient allocation of global capital [Claus et al. (2001)]. When an economy opens itself to private capital movements, the impact on investment depends on the environment of domestic investment and the objectives of investors [Amadou (2011)]. If the marginal returns on capital are higher than international interest rates, then substantial capital will enter the domestic economy and fund domestic investment. Blanchard and Giavazzi (2002), for example, found that the openness of Greece and Portugal, within the framework of their adhesion to European monetary union, lead to significant entries of capital which is used to finance investment. According to Kraay and Ventura (2000), foreign capital may still enter the domestic economy for portfolio diversification purposes. However, the overall strength of the relationship between domestic investment and foreign savings may be moderated by home bias, which in this respect suggests that there may be some barriers to capital mobility. A further consideration is that the availability of foreign funds depends, in part, on the sustainability of the domestic current account which, in turn, depends on the ability of the economy to generate sufficient trade surpluses in the future to repay existing debt and the willingness of foreign investors to continue lending [see, for example, Milesi-Ferretti and Razin (1996)].

The point of departure in the analysis of FH is the cross-section model:

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

where I is investment, S is domestic savings (both expressed as a percentage of GDP), u is the equation error term, and $i = 1, ..., N$ indicates the number of countries considered for estimation. In equation (1), the slope coefficient α_1, often referred to as the saving retention coefficient, measures the proportion of incremental savings that is invested domestically, and the intercept term is included to account for the fact that the series have non-zero means. While a zero slope coefficient is consistent with perfect capital mobility, a value of one would be supportive of no capital mobility. Within this framework, the basic estimates of α_1 originally obtained by FH strongly contradict the hypothesis of perfect world capital mobility. The conclusion reached by FH was irrespective of whether 15-year average data over the period 1960 to 1974 were used in the regression analysis or data for the 5-year sub-periods 1960-1964, 1965-1969 and 1970-1974; see also Sinn (1992) for similar findings based on year-by-year cross-section regressions.

Following the contribution of FH and the subsequent work by Feldstein (1983), a number of authors have also examined international capital mobility by employing an alternative avenue of research based on time-series estimation applied to data of specific countries, so that the model of interest becomes:

$$I_t = \beta_0 + \beta_1 S_t + u_t, \tag{2}$$

where the index $t = 1, ..., T$ denotes the number of time observations used for estimation. As expected, the growing availability of data over longer time spans across countries has led economists to use tools of the econometric analysis of non-stationary time series; see e.g. Coakley et al. (1998) for a survey of cross-section and time-series studies. Taylor (1996) highlights that the key distinction between equations (1) and (2) is that in the former the slope coefficient can be interpreted as a long-run saving retention coefficient, while in the latter it is to be interpreted as a short-run coefficient. Lastly, and perhaps not surprisingly, a natural extension that has also appeared in the literature relies on the use of panel data, in particular employing techniques developed for the case in which the time dimension, T, is much larger than the cross section dimension, N; see e.g. Ho (2002a, 2002b) and Adedeji and Thornton (2008).

The feature that is common across studies that have re-examined the FH approach to international capital mobility is that domestic investment is regressed on domestic savings; this is regardless of whether regression results are based on cross-section, time-series or panel data. In the estimating equations there is little or no explicit recognition that domestic investment may in fact be correlated with savings in one or more other countries. Moreover, this is a dimension of the analysis that is absent from the existing empirical contributions to the FH literature and is an issue that we now address in this study. Thus, following the existing literature, in this paper we begin by postulating the well-known relationship that states that in the presence of perfect capital immobility, domestic investment must be equal to domestic savings:

$$I_{i,t} = S_{i,t}, \tag{3}$$

where $i = 1, ..., N$ and $t = 1, ..., T$ denote the number of countries and time observations available for estimation, respectively. Then, we depart from the existing literature by multiplying equation (3) by minus one, and adding foreign savings at time t in both sides, which we denote $S_{j,t}$ where $j = 1, ..., N$ and $i \neq j$:

$$S_{j,t} - I_{i,t} = S_{j,t} - S_{i,t}. \tag{4}$$

This last equation can be alternatively expressed in terms of the following regression model:

$$S_{j,t} - I_{i,t} = \alpha_{0,ij} + \alpha_{1,ij}(S_{j,t} - S_{i,t}) + u_{ij,t}. \tag{5}$$

Equation (5), which can be estimated for each country pair (i, j) for all $i \neq j$, has characteristics that are worth describing in a little bit more detail. First, both the intercept and the slope coefficients are specific to each country pair (i, j), which means that we are implicitly assuming that the number of time observations available for estimation is sufficiently large, so that it is possible to estimate separate regressions for each country pair. Thus, the econometric relationship postulated in equation (5) can be viewed as an example of the type of regression models encountered in the so-called large T panel literature.

Second, putting aside the intercept term, as it simply accounts for units of measurement, and the error term, the economic interpretation of equation (5) is more general than that of the corresponding regression model in FH. Indeed, setting $\alpha_{1,ij} = 1$ in equation (5) results in the well-known relationship that states that domestic investment, $I_{i,t}$, must be equal to domestic savings, $S_{i,t}$, and therefore implies perfect capital immobility. Moreover, setting $\alpha_{1,ij} = 0$ is associated with perfect capital mobility since it results in domestic investment, $I_{i,t}$, moving in tandem with foreign savings, $S_{j,t}$.[1] In addition to this, we define *strong* capital mobility as where $0 < \alpha_{1,ij} < 0.5$ such that domestic investment is more closely correlated with foreign savings than with domestic savings.

Third, the relationship between domestic saving and investment expressed in equations (1) and (2) may be driven by factors such as productivity shocks, demographic changes, government policies, among others, which have little or nothing to do with capital mobility. Such endogeneity issues are less of a concern because equation (5) does not involve directly regressing domestic investment on domestic savings.

The empirical approach described above allows us to obtain estimates of the parameter of interest $\alpha_{1,ij}$ for all possible country pairs (i, j), where $i \neq j$.[2] This is clearly useful because it provides a measure of the extent of capital mobility that a country may have with respect to any other country in the world. This is in sharp contrast to the FH approach, where foreign savings are kept in the background as an aggregate variable. To put it in another way, when one applies the FH approach and fails to reject the null that

[1] We also include the possibility $\alpha_{1,ij} < 0$ as part of the case where capital is perfectly mobile. Since equation (5) can be rewritten as $I_{i,t} = -\alpha_{0,ij} + (1-\alpha_{1,ij})S_{j,t} + \alpha_{1,ij}S_{i,t} - u_{ij,t}$, $\alpha_{1,ij} < 0$ implies that $I_{i,t}$ is very sensitive to $S_{j,t}$ because $(1 - \alpha_{1,ij}) > 1$.

[2] The cases where $i = j$ are of no particular interest as they result in a regression of $(S_{i,t} - I_{i,t})$ against an intercept.

the saving retention coefficient is equal to zero, by default the alternative option is that domestic investment must therefore be surely being financed through foreign savings. In contrast, our employment of equation (5) sees the role played by foreign savings $S_{j,t}$ as explicitly incorporated. This approach also enables us to assess the extent of capital mobility through considering the percentage of cases whereby the null of perfect capital mobility is not rejected. It might be the case that the reliance on foreign savings increases, but it is of interest to know the extent to which this increased reliance is reflected in significant correlations involving a wider set of countries. Moreover, an increased acceptance rate of the zero null will imply that perfect capital mobility has become more broadly based involving a greater number of countries.

Lastly, Frankel (1992) points out that the FH definition of international capital mobility requires that the real interest rate in a country must be tied to the world real interest rate through the definition of real interest rate parity. Yet, existing empirical literature appears disarticulated, because studies of capital mobility have generally examined one definition independently from the other. By contrast, the pairwise approach adopted in this paper offers the advantage that one can explicitly incorporate the two definitions in one single regression equation, and test whether the likelihood of capital mobility between countries i and j (according to the FH definition) is related to the (absolute value of the) spread between their corresponding real interest rates. In other words, the smaller the spread between real interest rates, is it therefore more likely that there is increased capital mobility based on the FH definition?

In summary, our empirical modelling strategy involves two stages. In the first stage, equation (5) is estimated using time-series data for all possible country pairs (i, j), where $i = 1, ..., N$, $j = 1, ..., N$, $t = 1, ..., T$ and $i \neq j$. For all country pairs, we examine the case for perfect capital mobility by testing $H_0 : \alpha_{1,ij} = 0$ based on equation (5), and define the indicator variable

$ind_{ij} = 1$ if the null is not rejected (at some significance level), and zero otherwise. Next, in the second stage of the analysis, we move to a standard cross-section binary choice setup, where ind_{ij} is related to the absolute value of the difference between the real interest rates of countries i and j:

$$ind_{ij} = \delta_0 + \delta_1 |r_i - r_j| + \varepsilon_{ij}. \quad (6)$$

In equation (6), the condition postulated by Frankel (1992), where the validity of the FH definition of capital mobility is made dependent on the definition of real interest rate parity, can be easily addressed by testing $H_0 : \delta_1 = 0$, against the alternative that $H_1 : \delta_1 < 0$.

3 Data

We examine annual time series data, running from 1970 to 2011, on investment and gross savings for $N = 25$ OECD member countries, namely 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. The number of countries is thus larger than that originally studied by Feldstein and Horioka (1980) and Feldstein (1983).[3] Both investment and gross savings are expressed as a proportion of GDP. The data series were downloaded from the OECD iLibrary. At this point, one could well argue that given that a non-negligible part of international capital flows goes through non-OECD markets, it would also be interesting to include a sample of emerging markets. Thus, in addition to the 25 countries listed above, we also considered an augmented sample of 38 countries that includes the following countries: Argentina, Brazil, Chile, Colombia, India,

[3]Feldstein and Horioka (1980) do not include in their estimations France, Iceland, Luxembourg, Mexico, Norway, Portugal, Spain, Switzerland and Turkey. In addition, West Germany in the original FH sample is replaced by Germany.

Israel, Korea, Malaysia, Morocco, Pakistan, Peru, Philippines and Thailand. For these additional 13 countries the investment and gross savings data (as a percentage of GDP) are obtained from the World Bank website.[4] The selection of the countries and of the sample period is guided by the need to collect a balanced data panel for the purposes of the econometric analysis.

4 Time-series properties of the data

We are first interested in determining whether innovations (or shocks) to the investment and savings series (as a percentage to GDP) are cross-sectionally independent, as well as in establishing the order of integration of these macroeconomic aggregates.

Starting with an examination of cross sectional independence, we perform the Pesaran (2004) test which involves estimating augmented Dickey and Fuller (1979) type regressions to each country variable, using p lags of the dependent variable, and denoting the resulting residuals of the individual series as $\hat{e}_{it}$ (the purpose of this initial stage is to capture any serial correlation that is present in the underlying series). Then, the cross-correlation coefficient between the residuals of cross sections i and j is computed as:

$$\hat{\rho}_{ij} = \frac{\sum_{t=1}^{T} \hat{e}_{it}\hat{e}_{jt}}{\left(\sum_{t=1}^{T} \hat{e}_{it}^2\right)^{1/2} \left(\sum_{t=1}^{T} \hat{e}_{jt}^2\right)^{1/2}}. \tag{7}$$

Lastly, the CD statistic is calculated as:

$$CD = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{ij} \right) \sim N(0,1). \tag{8}$$

Pesaran (2004) shows that under the null hypothesis of cross sectional independence, the statistic given in equation (8) follows a standard normal

[4] Chile, Israel and Korea entered the OECD in more recent years and for this reason they are not included in our group of 25 countries. For these three countries the source of the data is also the World Bank.

distribution. Furthermore, Monte Carlo simulation experiments presented by Pesaran indicate that the CD test exhibits correct size and satisfactory power in small samples, being also robust to the presence of structural breaks and unit roots. Table 1 summarises the results of applying the CD statistic to the 38 investment and savings ratios, when intercept and trend are included in the deterministic part of the test regressions, and $p = 6$ lags of the dependent variable are used in the ADF-type regressions.[5] According to the results, the null of cross sectional independence is clearly rejected for both the investment and savings ratios. A qualitatively similar conclusion is reached when other lag orders are considered, or when an intercept is the only deterministic component included in the test regressions.

The order of integration of the investment and savings data series is examined using the Hadri (2000) panel stationarity test, which is the panel version of the well-known univariate Kwiatkowski et al. (1992) (KPSS) stationarity test around a mean (trend). The Hadri test, which tests the joint null hypothesis that all individual series in a panel are mean (trend) stationary against the alternative that at least one of them is not, offers the advantage that failure to reject the null implies that all the macroeconomic series in the panel are stationary. To calculate the Hadri test we start by computing the KPSS statistic, where the optimal number of lags required to account for serial correlation is determined by a General-To-Specific (GETS) algorithm. GETS involves fitting an AR(p) model to the de-meaned (detrended) series, and checking 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. 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). Monte

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

Carlo experiments undertaken by Carrión-i-Silvestre and Sansó (2006) indicate that this method renders a statistic with correct size and good power properties. In addition to this, we apply an autoregressive bootstrap to the Hadri statistic in order to account for cross sectional dependence (which we already know is a feature that is present in the investment and saving series under analysis); see e.g. Hadri and Rao (2008) for details on the implementation of the bootstrap procedure. To the extent that the Hadri test employed in this paper accounts for cross sectional dependence through the bootstrap methodology, the test belongs to the so-called family of second generation panel unit root and stationarity tests, following the terminology introduced by Breitung and Pesaran (2008).[6]

Table 2 presents the results of the KPSS and Hadri tests for stationarity around a trend. Looking first at the outcome of the individual KPSS results, in the overwhelming majority of the cases the series appear to be stationary, with 32 out of 38 in the case of investment, and 31 out of 38 in the case of savings. The finding of stationarity is confirmed when the series are considered as a data panel, and after accounting for the presence of cross sectional dependence through a bootstrap procedure. Indeed, according to the results in the last row of Table 2, we fail to reject the null hypotheses of trend stationarity hypothesis for both the investment and savings series (with respective p-values of 0.988 and 0.628). The finding of stationarity clearly rules out the possibility of an analysis of the investment and saving series within a cointegration framework, but rather using standard tools from regression analysis.[7]

[6]Given the extension of the time period it is also worth considering the effect of possible structural breaks. For this, we implemented the more recent Hadri and Rao (2008) test, which tests the null of joint stationarity allowing for the possibility of one-time endogenously determined structural breaks as well as cross section dependence. Results not reported here indicate that the structural breaks identified by the Hadri and Rao testing procedure are not sufficient to alter our conclusion regarding the stationarity of the investment and savings series over the period of analysis.

[7]Ericsson et al. (1994) and Juselius (1999) argue that the order of integration is not a property inherent to a time series. That is, a time series may be integrated of order one during a specific sample period, and integrated of order zero during another one.

Interestingly, had we incorrectly assumed cross-sectional independence, and therefore compared the calculated Hadri statistic against critical values derived from a standard normal distribution, the null hypothesis would then be incorrectly rejected.

5 A pairwise test of the FH definition of capital mobility

Ordinary least squares estimation of equation (5) for all possible country pairs (i, j), where $i = 1, ..., N$, $j = 1, ..., N$ and $i \neq j$ are reported in Table 3. Given that in the panel of $N = 25$ ($N = 38$) countries the total number of estimated regressions is 600 (1406), and the resulting number of estimated coefficients is very large, it is possibly better to base our comparisons on the average value of $\hat{\alpha}_{1,ij}$, which we denote as $\overline{\hat{\alpha}}_{1,ij}$. Here, we are also particularly interested in testing the null of perfect capital mobility which is given by $\alpha_{1,ij} = 0$. Estimation was initially performed over the whole 1970-2011 sample period. However, we also assess the time-varying nature of capital mobility over changing sample periods using a moving window that spans a period of 22 years. Thus, for the rolling regressions we start with the sample sub-period 1970-1991, then 1971-1992, and so on until 1990-2011. The reason for choosing a 22-year window size in the rolling regression will become apparent in the next section.

The results reported in the top row of Table 3 indicate an average value of $\hat{\alpha}_{1,ij}$ of 0.612 over the full sample period for the panel of 25 OECD countries. Further, in 83 out of the 600 estimated regressions (that is, 13.8% of the cases) we fail to reject the null of perfect capital mobility at a 10% significance level. In turn, the results of the larger panel of 38 countries yield an average value of $\hat{\alpha}_{1,ij}$ of 0.656 which is more supportive of capital immobility. Furthermore, the percentage of cases where we fail to reject the null of perfect capital mobility is lower when the larger panel including emerging economies is used. Thus, it appears that with the inclusion of emerging markets the evidence in favour of capital mobility is somehow weaker. As noted by Chang and Smith (2014), it is puzzling that the saving-investment correlations for emerging economies found in existing studies of the FH hypothesis have tended to be significantly lower than for advanced economies, though still positive. In contrast, our findings imply that capital mobility involving emerging economies is actually lower thereby offering to resolve the puzzle observed on the basis of earlier work.

Turning to the results of the rolling regressions, the average value of $\hat{\alpha}_{1,ij}$ which can be taken as a measure of depth of capital mobility starts to fall towards a value of 0.593 over the period 1990-2011. This period of increased correlation with foreign savings corresponds to the introduction of the Euro single currency which has a direct relevance for the twelve joining members of our sample plus Denmark which has remained outside of the single currency but has maintained close ties with Euro members. In addition to this, the percentage of times we fail to reject the null of capital mobility has been increasing consistently over time, passing from 13% during 1970-1991 to 21.5% during 1990-2011. As to the larger panel of 38 countries, the upward movement in the percentage of times we fail to reject the null of capital mobility is observed starting from the sub-period 1976-1997, passing from 8.8% to 18.6% during 1990-2011. These findings are consistent with foreign savings being sourced from a wider set of countries. Overall, these results are thus supportive of an increased extent of capital mobility over the years which is a finding documented elsewhere; see, for example, Adedeji and Thornton (2008). If we look at the years that immediately follow the global financial crisis, the percentage of cases where we fail to reject the null of perfect capital mobility is greatest for the OECD panel thereby indicating weaker capital mobility involving the emerging economies.

However, focussing on the panel for 25 countries we can qualify this somewhat and argue that the depth of capital mobility is weak or limited. This is because $\overline{\hat{\alpha}}_{1,ij}$ is never below 0.5. This might suggest that domestic over foreign savings have on average still been the more important consideration in explaining domestic investment rates where domestic agents are still more likely to invest locally than globally. This is consistent with home bias where the international diversification of portfolios is constrained by a range of transactions and information acquisition costs and uncertainties associated with conducting business in a foreign market; see Georgopoulos and Hejazi (2005), for example. Although we find that home bias is present in increasingly integrated global financial markets, the degree of home bias has generally fallen. Further to this, the majority of pairwise cases (78.5%) is still characterised by the null of perfect capital mobility being rejected in the final moving window. Indeed, there have been short-run episodes (namely 1997, 2003-4 and 2010) where the percentage of rejections has actually increased thereby running counter the general increase in capital mobility over time. We find that 223 pairs out of 600 (only 37% of the sample) are characterised by $\hat{\alpha}_{1,ij} < 0.5$ such that domestic investment is more strongly correlated with foreign rather than domestic savings. With regards to the most recent decline in capital mobility, one can note the work by Ahrend and Schwellnus (2012) who point out that the global crisis of 2008-09 went in hand with sharp fluctuations in capital flows with uncertainty-averse investors then selling assets for which they had poor information, including those in geographically-distant regions.

Thus far, the results are focused on the point estimate of $\overline{\hat{\alpha}}_{1,ij}$. However, it is also important to consider the precision of these estimates because the presence of cross-section dependence is likely to increase the uncertainty. We address the cross section dependence that characterises the macroeconomic aggregates being studied through adapting the factor augmented sieve bootstrap approach described in Pesaran et al. (2009). In this approach, global shocks can induce cross-section dependence which is interpreted in terms of an unobserved common factor model (see Appendix A for a description of the steps involved to implement the bootstrap). The parameters of the underlying factor model, which is treated as an approximation to the true data generation process, are estimated directly, and subsequently used to obtain bootstrap samples of the investment (saving) share on GDP for each country (the number of bootstrap replications is set equal to $B = 10000$). Lastly, these bootstrap samples are used to estimate equation (5). In what follows, we shall compare the results of the most recent sample period 1990-2011 with those obtained for 1978-1999, right before the start of the Euro single currency area, and also the period during which the highest point estimate of $\bar{\hat{\alpha}}_{1,ij}$ is observed. Our results indicate that the mean of the bootstrap distributions of $\bar{\hat{\alpha}}_{1,ij}$ are equal to 0.615 and 0.565 during 1978-1999 and 1990-2011, respectively. These bootstrap estimates are below the corresponding point estimates reported in Table 3, that is 0.686 and 0.593, respectively. The 90% bootstrap confidence intervals, which range from 0.486 to 0.745 for 1978-1999 and from 0.481 to 0.645 for 1990-2011, do cover the corresponding point estimates as well as 0.5. The number of times we fail to reject the null of capital mobility has increased between 1999 and 2011, and both confidence intervals around $\bar{\hat{\alpha}}_{1,ij}$ still embody 0.5 which gives rise to the possibility of strong capital mobility with domestic investment more correlated with foreign rather than domestic savings. With no evidence of a significant increase in $\bar{\hat{\alpha}}_{1,ij}$, there is less clear evidence that capital mobility has increased since the beginning of the single currency era in 1999.

Our pair-wise results point towards a limited depth and extent of capital mobility that has been increasing over time. This is the case for the two samples of countries that we consider. During the study period, the sample of 25 OECD countries has comprised 12 countries that have joined the Euro single currency (Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain). A priori, we would expect the pairings based only on Euro members to exhibit a much greater extent of capital mobility between them than is the case for other pairings on account of the absence of capital controls. In terms of existing work based on the FH hypothesis, Kapapoulos and Paleologos (2004) find that domestic saving and investment within Eurozone is less correlated than they were before the advent of the Euro. By contrast, Alam and Islam (2010) conclude that there is strong evidence in favour of higher degree of capital mobility for many EU countries, especially for the Eurozone countries following the introduction of the single currency. Bereau (2007), on the other hand, finds that the introduction of the Euro has not significantly accelerated the process towards a high degree of financial integration in Europe. For the purpose of our investigation, the 12 Euro members constitute 132 Euro-only pairs out of the full OECD sample of 600 pairs (22%) while the remaining 468 pairs comprise at least one non-Euro country. In a further investigation into the depth and extent of capital mobility, we find that for the 1990-2011 period, 56 out of 132 (42%) of the Euro-only pairs are characterised by $\hat{\alpha}_{1,ij} < 0.5$. In other words, a greater proportion of the Euro-only sub-sample is characterised by domestic investment being more strongly correlated with foreign rather than domestic savings. This compares with only 167 out of 468 (36%) of the remaining pairs comprising at least one non-Euro country having $\hat{\alpha}_{1,ij} < 0.5$ (see Figure 1). That is, the majority of the non-Euro sub-sample is characterised by domestic investment being more strongly correlated with domestic rather than foreign savings.[8]

6 The FH definition of capital mobility and interest rate differentials

In this section we test the argument made by Frankel (1992) that the FH definition of capital mobility requires the validity of real interest rate parity,

[8]If we instead consider an EU-only sub-sample of 15 countries by adding Denmark, Sweden and the UK to the 12 Euro members, then 85 out of 210 (40%) of the EU-only pairs are characterised by $\hat{\alpha}_{1,ij} < 0.5$. While the extent of capital mobility is slightly lower than in the Euro-only case, it is nonetheless greater than for the rest of the sample.

which was formalised earlier in equation (6). To do this, we consider the estimated values of $\hat{\alpha}_{1,ij}$ (and associated standard errors) that were obtained using the data for the sub-period 1990-2011 to define the indicator variable ind_{ij}, which takes the value of one if the null of perfect capital mobility is not rejected at the 10% significance level, and zero otherwise; this indicator variable also takes the value of one when $\hat{\alpha}_{1,ij} < 0$ and also statistically significant, because of the reason exposed in footnote 1. The question that underlies the following analysis is whether there is an association between the indicator variable ind_{ij} and the real interest rate differential between countries i and j. Bearing in mind the need to keep the data across countries as consistent as possible, we assembled a full set of real interest rates for the sub-period 1990-2011 based on OECD data, corresponding to the 3-month nominal interest rate minus (consumer price index) inflation, with only a few proxies drawn on World Bank data. Unfortunately, a complete data set based on the full 1970-2011 sample period is problematic for some of the countries included in the analysis, where gaps prior to 1990 are mostly on account of missing consistent nominal interest rate data.

Employing average real interest rate data for each of the 25 OECD countries calculated over the same sub-period 1990-2011, estimation of the binary regression equation postulated in equation (6) using a probit model yields:

$$\begin{aligned} ind_{ij} \;&=\; \underset{(0.098)}{-\,0.610} \;-\; \underset{(0.078)}{0.172}\;|r_i - r_j|, \\ &\text{Pseudo}-\text{R}^2 = 0.008,\ \text{Obs.} = 600, \end{aligned} \tag{9}$$

where standard errors are reported in parentheses. As can be seen, the estimated slope coefficient in the probit model has the expected negative sign and is statistically different from zero pointing to an increased probability of perfect capital mobility between the bivariate pairs as the absolute real interest rate differential becomes smaller.[9] The marginal effect, calculated at

[9]When the indicator variable is defined as taking the value of one when the null of capital mobility is not rejected at the 5% (1%) significance level, the estimated slope

the mean of the regressor over the estimation range, is equal to -0.050. This result suggests that increased capital mobility (as implied by the FH definition of capital mobility) is associated with a reduced absolute real interest differential. This would appear to lend support to the consistency of the FH hypothesis (a quantity measure) compared with other ways of judging financial integration (a price measure).

Estimation of the previous binary-choice model specification using the extended sample of 38 countries yields the following results:

$$\begin{aligned} ind_{ij} &= \underset{(0.045)}{-\,0.772} - \underset{(0.006)}{0.029}\ |r_i - r_j|, \\ &\text{Pseudo}-\text{R}^2 = 0.022,\ \text{Obs.} = 1406, \end{aligned} \tag{10}$$

Once again, the estimate of the slope coefficient has the expected negative sign and is statistically different from zero. Although the marginal effect is now -0.007 which, in absolute terms, is smaller than the value obtained when using the 25 OECD country data, this finding nonetheless lends support to the idea that the FH definition of capital mobility requires the validity of real interest rate parity.

7 Concluding remarks

Despite increases in capital mobility over the years, a puzzle remains as to why correlations between domestic investment and domestic savings have remained high thereby in conflict with the Feldstein-Horioka approach. In this paper, we have proposed an alternative way of testing for capital mobility based on a pairwise procedure that explicitly incorporates the potential correlation between domestic investment and foreign savings.

Our results using a panel of OECD and emerging market economies initially suggests that the depth and extent of capital mobility remains generally

coefficient in equation (6) remains negative although the resulting model is somewhat inferior and therefore not reported here.

limited, and that mobility has increased over the past twenty years. However, stronger support for capital mobility is found among the single currency Euro-only pairs. Furthermore, our approach indicates that saving-investment correlations involving emerging market economies are greater than between developed market economies. This latter finding perhaps resolves an existing puzzle insofar as studies have found that saving-investment correlations for emerging economies tend to be significantly lower than for advanced economies. Either way, limited capital mobility may be due to home country bias in portfolio formation. If one expects that a diversified portfolio will result in higher expected profits under the same risk, then one is led to ask about the extent to which international portfolios are optimally composed. Indeed, one might expect countries to smooth savings and investment expenditure over time through rebalancing their portfolios using foreign assets as a buffer stock. Based on our findings, this is most likely present among Euro members and least likely among emerging market economies.

In terms of policy implications that arise from our study, there is a case that many countries could effectively adopt policies that focus on increasing investment through increasing domestic savings. For this purpose, fiscal policy could be geared towards microeconomic incentives that impact on domestic household behaviour. The need for such policy is perhaps stronger in the case of emerging market countries that are relatively more reliant on domestic savings as a means of funding investment. However, for those countries that adopt a single currency, then an emphasis on focusing policy on domestic savings in order to stimulate investment is perhaps less strong. Instead, fiscal policy may be conducted at a more international level aimed at raising the general pool of savings. This provides more focus on the debate concerning the desirability of fiscal integration across member states.

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Table 1: CD cross-section independence test

Panel	CD test	p-value
Investment	15.451	[0.000]
Savings	22.793	[0.000]

Note: 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.

Table 2: Trend-stationarity tests

	Investment		Savings	
Country	Lag	KPSS	Lag	KPSS
Argentina	6	0.185‡	2	0.071
Australia	1	0.058	1	0.077
Austria	1	0.043	1	0.082
Belgium	2	0.085	5	0.290‡
Brazil	2	0.075	1	0.060
Canada	2	0.085	1	0.041
Chile	1	0.058	3	0.057
Colombia	3	0.084	1	0.019
Denmark	2	0.114	6	0.242‡
Finland	2	0.096	2	0.084
France	2	0.111	1	0.059
Germany	2	0.125	1	0.093
Greece	2	0.106	1	0.102
Iceland	5	0.045	1	0.065
India	4	0.122	1	0.078
Ireland	6	0.140	6	0.173†
Israel	5	0.221‡	1	0.100
Italy	2	0.111	1	0.036
Japan	6	0.115	1	0.028
Korea	6	0.293‡	1	0.082
Luxembourg	1	0.052	1	0.124
Malaysia	2	0.098	1	0.060
Mexico	2	0.099	1	0.030
Morocco	4	0.075	6	0.325‡
Netherlands	2	0.117	1	0.045
New Zealand	2	0.077	6	0.210†
Norway	2	0.079	1	0.091
Pakistan	2	0.111	1	0.049
Peru	1	0.063	1	0.037
Philippines	4	0.139	1	0.065
Portugal	6	0.180†	2	0.115
Spain	2	0.087	2	0.089
Sweden	2	0.078	6	0.418‡
Switzerland	2	0.101	5	0.382‡
Thailand	2	0.104	5	0.085
Turkey	1	0.072	1	0.060
United Kingdom	6	0.283‡	2	0.116
United States	6	0.180†	6	0.149
Hadri test		6.502		6.601
Bootstrap p-value		[0.988]		[0.628]

Note: † and ‡ indicate 5% and 1% significance levels based on Sephton (1995). The p-values of the Hadri test are based on 2000 bootstrap replications.

Table 3: Pairwise test of the FH definition of capital mobility

		Countries = 25		Countries = 38	
Start	End	$\overline{\hat{\alpha}}_{1,ij}$	Capital mobility	$\overline{\hat{\alpha}}_{1,ij}$	Capital mobility
1970	2011	0.612	0.138	0.656	0.122
1970	1991	0.667	0.130	0.681	0.146
1971	1992	0.671	0.132	0.684	0.142
1972	1993	0.673	0.137	0.683	0.124
1973	1994	0.670	0.137	0.676	0.107
1974	1995	0.668	0.147	0.684	0.111
1975	1996	0.676	0.150	0.692	0.111
1976	1997	0.686	0.138	0.696	0.088
1977	1998	0.685	0.138	0.685	0.107
1978	1999	0.686	0.143	0.685	0.164
1979	2000	0.671	0.163	0.677	0.172
1980	2001	0.668	0.185	0.682	0.179
1981	2002	0.662	0.197	0.680	0.188
1982	2003	0.667	0.178	0.680	0.180
1983	2004	0.658	0.170	0.672	0.195
1984	2005	0.640	0.183	0.654	0.202
1985	2006	0.620	0.188	0.634	0.223
1986	2007	0.603	0.198	0.621	0.213
1987	2008	0.600	0.203	0.614	0.216
1988	2009	0.595	0.223	0.626	0.225
1989	2010	0.593	0.208	0.631	0.199
1990	2011	0.593	0.215	0.625	0.186

Note: $\overline{\hat{\alpha}}_{1,ij}$ refers to the simple average of $\hat{\alpha}_{1,ij}$, which is calculated over all possible country pairs (i, j), where $i = 1, ..., N$, $j = 1, ..., N$ and $i \neq j$. For 25 (38) countries the total number of pairs under consideration is 600 (1406). Capital mobility refers to the percentage of times we fail to reject the null hypothesis of perfect capital mobility, that is $H_0 : \alpha_{1,ij} = 0$, at the 10% significance level.

A Appendix

Bootstrapping the empirical distribution of $\overline{\hat{\alpha}}_{1,ij}$

To bootstrap $\overline{\hat{\alpha}}_{1,ij}$ we follow Pesaran et al. (2009) who consider the following model set up:

$$y_{it} = \boldsymbol{\delta}_i^{'}\mathbf{d}_t + \boldsymbol{\gamma}_i^{'}\mathbf{f}_t + \varepsilon_{it} \tag{A.1}$$

$$\Delta\varepsilon_{it} = \eta_i + \lambda_i\varepsilon_{i,t-1} + \sum_{l=1}^{p_i}\psi_{il}\Delta\varepsilon_{i,t-l} + \upsilon_{it} \tag{A.2}$$

$$\Delta f_{st} = \boldsymbol{\mu}_i^{'}\mathbf{d}_t + \phi_s f_{s,t-1} + \sum_{l=1}^{p_s}\xi_{sl}\Delta f_{s,t-l} + e_{st} \tag{A.3}$$

where y_{it} is investment (saving) as a proportion of GDP in country i at time t, $s = 1, 2, ..., m$ is the number of common factors, $\mathbf{d}_t = (1, t)^{'}$ is a vector of deterministic components that includes intercept and trend, $\mathbf{f}_t$ is a $(m \times 1)$ vector of unobserved common factors, with elements denoted f_{st}, and ε_{it} denotes the corresponding idiosyncratic elements. The unobserved common factors f_{st} and (or) the idiosyncratic elements ε_{it} may be $I(0)$ or $I(1)$.

In line with Pesaran et al. (2009), we use the cross-sectional average of y_{it}, denoted $\bar{y}_t = N^{-1}\sum_{i=1}^{N} y_{it}$, as an estimate of the common factor that may induce cross-section dependence across countries. To account for cross-section dependence, investment (saving) in each country is regressed on $\bar{y}_t$:

$$y_{it} = \hat{\delta}_{1i} + \hat{\delta}_{2i}t + \hat{\gamma}_{1i}\bar{y}_t + \hat{\varepsilon}_{it}. \tag{A.4}$$

It should be noted that in these factor equations the trend term is included if it is statistically significant at the 5% level.

The next step is to examine the time-series properties of $\bar{y}_t$, which may be $I(0)$ or $I(1)$. This involves estimating the ADF(p) regression:

$$\Delta\bar{y}_t = \hat{\mu} + \hat{\phi}\bar{y}_{t-1} + \sum_{l=1}^{p}\hat{b}_l\Delta\bar{y}_{t-l} + \hat{e}_t, \tag{A.5}$$

which may also include trend if it is statistically significant, and where the number of lags p may be determined e.g. using the Akaike information criterion (AIC). To illustrate the implementation of the bootstrap, let us consider the case in which $\bar{y}_t$ does not have a unit root, nor a deterministic trend (which appear to be the features that best characterise the resulting cross-sectional means of the investment and saving series). In this case, the bootstrap samples of $\bar{y}_t$, denoted $\bar{y}_t^{(b)}$, can be computed using the following generating mechanism:

$$\bar{y}_t^{(b)} = \hat{\mu} + (1 + \hat{\phi})\bar{y}_{t-1}^{(b)} + \sum_{l=1}^{p}\hat{b}_l\Delta\bar{y}_{t-l}^{(b)} + \hat{e}_t^{(b)} \tag{A.6}$$

where bootstrap residuals $\hat{e}_t^{(b)}$ are generated by randomly drawing with replacement from the set of estimated and centred residuals $\hat{e}_t$ in equation (A.5), and where the first $(p+1)$ values of $\bar{y}_t$ are used to initialise the process $\bar{y}_t^{(b)}$.

In turn, the bootstrap samples of y_{it}, denoted $y_{it}^{(b)}$, are generated as:

$$y_{it}^{(b)} = \hat{\delta}_{1i} + \hat{\delta}_{2i}t + \hat{\gamma}_{1i}\bar{y}_t^{(b)} + \hat{\varepsilon}_{it}^{(b)}, \tag{A.7}$$

where $\hat{\delta}_{1i}$, $\hat{\delta}_{2i}$ and $\hat{\gamma}_{1i}$ are the OLS estimates of δ_{1i}, δ_{2i} and γ_{1i} in equation (A.4), respectively, and

$$\varepsilon_{it}^{(b)} = \hat{\eta}_i + (1+\hat{\lambda}_i)\varepsilon_{i,t-1}^{(b)} + \sum_{l=1}^{p_i} \hat{\psi}_{il}\Delta\varepsilon_{i,t-l}^{(b)} + \upsilon_{it}^{(b)}, \tag{A.8}$$

where bootstrap residuals $\upsilon_{it}^{(b)}$ are generated by randomly drawing with replacement from the set of estimated residuals υ_{it} in equation (A.2), and where the first $(p+1)$ values of $\hat{\varepsilon}_{it}$ are used to initialise the process $\hat{\varepsilon}_{it}^{(b)}$. The AIC is used to determine the optimal number of lags.

The bootstrap procedure described above is applied to obtain the bootstrap samples of investment and saving (both as a share of GDP) by setting $y_{it} = I_{it}$ and $y_{it} = S_{it}$, respectively. Letting $I_{it}^{(b)}$ and $S_{it}^{(b)}$ be the respective bootstrap samples of I_{it} and S_{it}, equation (5) is estimated $b = 1, ..., B$ times to derive the empirical distribution of $\overline{\hat{\alpha}}_{1,ij}$.

Figure 1: Empirical cumulative distribution functions of $\hat{\alpha}_{1,ij}$: 1990-2011