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Abstract

A systems *GMM* estimation method is used to estimate the Feldstein-Horioka equation from 1960-2007 with a panel of 12 OECD countries. It is found that the Feldstein-Horioka puzzle exists in a weaker form with a much reduced saving retention coefficient. The Bretton Woods agreement in particular has weakened the Feldstein-Horioka puzzle by significantly improving international capital mobility. In comparison the Maastricht agreement seems to have improved capital mobility only by a small magnitude. The Blundell and Bond approach systems *GMM* method and the structural break tests of Mancini-Griffoli and Pauwels are used in this paper.

Keywords: Feldstein-Horioka puzzle**,** Structural breaks, Effects of Bretton Woods and Maastricht agreements on international capital mobility.

JEL: C23, F21, F36

1. Introduction

The high correlation between domestic savings and investment is a stylized fact. Well known as the Feldstein-Horioka puzzle (henceforth FHP) it started with the seminal work of Feldstein and Horioka (1980, henceforth FH). They empirically showed that in a cross-section consisting of 16 OECD countries for the period 1960-1974, investment and saving are highly correlated, and argued that this provides evidence against international capital mobility. FH reasoned that saving and investment should be unrelated in an open economy since savings seek higher global returns. Capital mobility is important because it has implications for single currency debates, tax policies on capital and saving, whether growth is constrained by domestic saving rate and if fiscal deficits will have large crowding out effects on private investment. On the other hand if capital mobility is high, countries cannot pursue independent monetary policies. Because of these important policy implications Obstfeld and Rogoff (2000) and Sinha and Sinha (2004) have called FHP the mother of all puzzles. This puzzle, in spite of a number of empirical investigations with alternative specifications and estimation technique, still remains a puzzle. Recently the vast empirical literature on FHP is comprehensively surveyed by Apergis and Tsoumas (2009). They conclude that the majority of the empirical studies oppose the original strong results of FH but found that this correlation still exists in a weaker form. Furthermore, Apergis and Tsoumas take the view that the results in these studies are difficult to analyze beyond any doubt.

In light of the above observations it would be foolhardy to claim that our present paper is the final nail in the coffin of FHP. Our objective is to fill a gap in the existing results based on a number of alternative estimation methods. Apergis and Tsoumas draw attention in particular to some methodological differences in estimating the FH equation with the levels of the variables or with their first differences using panel data methods. However, it is possible to estimate both with the levels and first differences of the variables with panel methods of Blundell and Bond (1998) and also use the structural break tests of Mancini-Griffoli and Pauwels (2006). We shall discuss the merits of these two developments later in the paper.

The outline of this paper is as follows. Section 2 briefly reviews a few relevant empirical works and summarizes their main points in a table. Section 3 explains the Blundell and Bond approach and the structural break tests of Mancini-Griffoli and Pauwels. Empirical results are presented in

Section 4 with panel data for 13 OECD countries for the period 1960 to 2007 and Section 5 concludes.

2. Survey of Empirical Literature

Existing empirical studies on FHP have used cross sectional, panel data and time series methods for estimation. From our paper's perspective studies that have made significant contributions to the literature on the OECD countries are Feldstein and Horioka (1980), Feldstein (1983), Sachs (1981, 1983), Caprio and Howard (1984), Penati and Dooley (1984), Feldstein and Bachetta (1991), Tesar (1991), Bodman (1995), Coakley *et.al* (1996, 2003 and 2004), Ghosh (1995), Lemmen and Eijffinger (1995), Barkoulas *et al.* (1996), Apergis and Tsoulfidis (1997), Hussein (1998), Kim (2001), Blanchard and Giavazzi (2002), Amirkhalkhali *et al.* (2003), Kasuga (2004), Giannone and Lenza (2004), Bahmani-Oskooee and Chakrabarti (2005), Georgopoulos and Hejazi (2005), Chakrabarti (2006), Katsimi and Moutos (2007), Di Iorio and Fachin (2007), Christopoulos (2007), Grier *et.al* (2008) and Fouquau *et.al* (2009). For convenience the major findings of some from these studies are tabulated in Table 1. Essentially these studies estimate a simple equation of the following form or its variants:

$$ITY_{it} = \alpha_i + \beta_i STY_{it} + \varepsilon_{it} \qquad (1)$$

where ITY = ratio of investment to income and STY = ratio of saving to income, i and t are country and time subscripts and $\varepsilon_{it} \sim N(0,\sigma)$ for all i and t. The controversy is on the estimate of β, known as the saving retention coefficient. Under complete capital mobility β should be near zero. FH interpret this coefficient as an indicator of the degree of international capital mobility. Table 1 provides the estimated values of β in some key empirical works.

Table 1. Studies of FH Puzzle and their Findings

Authors	Period	Country	Methodology	Estimate of β	Major Findings
Feldstein and Horioka (1980)	1960-1974	16 OECD	Cross section	0.85 to 0.95	Low capital mobility exists.
Tesar (1991)	1960-1986 1960-1974 1975-1986	23 OECD	Cross section	0.840 0.870 0.810	Low capital mobility exists.
Coakley et.al (2001)	1980Q1 to 2000Q4	12 OECD	Panel Mean Group	0.32	Supports long run capital mobility and integration of international financial markets.
Giannone and Lenza (2004)	1970-1999 1970-1979 1980-1989 1990-1999 1980-1999	24 OECD	FAPR	0.34 0.50 0.21 0.22 0.18	Increased capital mobility in international financial markets.
Fouquau et.al (2009)	1960-2000	24 OECD	PSTR	0.710 0.704 0.526	Strong heterogeneity in the degree of mobility of OECD countries.
Katsimi and Moutos (2007)	1986-2002 1986-1990 1991-1995 1996-2000 1997-2002	25 OECD	OLS	0.572 0.611 0.702 0.372 0.261	Adding human capital investment does not alter the Feldstein and Horioka (1980) results significantly.
Di Iorio and Fachin (2007)	1960-2002	12 EU	FMOLS	from 0.590 to 1.030	The bootstrap panel stability tests confirmed cointegration with at least one break.

continued

Authors	Period	Country	Methodology	Estimate of β	Major Findings
Grier et.al (2008)	1947Q1- 2007Q1	USA	Bai-Perron (1998, 2003)	-	The saving and investment rates are stationary and not linked in the long run. Saving rate has two structural breaks in its mean and the investment rate is without a break.

Notes: The reported estimate of β in Fouquau et.al (2009) are based on adding 3 additional variables to equation (1) viz., degree of openness, size of the country and ratio of current account to GDP. For Giannone and Lenza (2004), we only report the estimates based on two common factors estimated with principal components. FAPR and PSTR is factor augmented panel regression and panel smooth threshold regression model, respectively. OLS, FMOLS, DOLS means Ordinary Least Squares, Fully Modified Ordinary Least Squares and Dynamic Ordinary Least Squares.

The pioneering work of Feldstein and Horioka (1980) showed that in a cross section consisting of 16 OECD countries for the period 1960-1974, β is close to unity, ranging from 0.85 to 0.95, in all cases. The low capital mobility persisted even when the degree of a country's openness or its size is taken into account. The original FH findings were confirmed by Feldstein (1983) and Feldstein and Bachetta (1991) by extending the sample period to 1960–1979 and 1960-1986, respectively which include observations from the post Bretton Woods agreement. These works found that β had not changed significantly.

Tesar (1991) used net savings and investment rates to estimate the FH equation for 23 OECD countries. This is a minor improvement since it is hard to estimate net investment and savings data because depreciation is an accounting concept and generally assumed a constant. His cross section estimate of the savings retention coefficient β is around 0.8 to 0.9 for the whole sample 1960-1986 and sub-sample 1960-1974 and 1975-1986 periods. By and large his results confirm FH's original estimates although Tesar's estimates of β are marginally less in the post Bretton Woods sample. Coakley et.al (2001) found that savings and investment are unit root variables and used time series panel data methods for estimation for 12 OECD countries for the period

1980Q1-2000Q4. Their estimates of β are much less at around 0.32 and support long run capital mobility and the integration of international financial markets.

Giannone and Lenza (2004) have used the Factor Augmented Panel Regression (FAPR) technique to estimate β for 24 OECD countries for the period 1970-1999. This approach allows for heterogeneous response of savings and investment to global shocks. Their results show that the homogeneity restriction on the propagation of global shocks across countries is rejected by the data. When the homogeneity assumption is relaxed, estimates of β reduced to 0.18 in the sample for 1990-1999. Recently, Fouquau et.al (2009) evaluated the FHP using Panel Smooth Threshold Regression Model (PSTR), developed by Gonzalez et al. (2005), to estimate β for 24 OECD countries for the period 1960-2000. While the country specific βs vary largely, their panel based estimates range between 0.5 to 0.7. They found that savings and investment relation is non-linear and the degree of openness, the size of the country and the ratio of current account balance to GDP have significant effects on the estimates of β.[1] Katsimi and Moutos (2007) have investigated whether ignoring investment in human capital has a significant effect on the estimate of β. In their sample of 25 OECD countries for the period 1986-2002 they found that estimates of β range from 0.572 for full sample to a low of 0.261 for the period 1997-2002.

For the purpose of testing breaks in the cointegrated panels, Di Iorio and Fachin (2007) have used panel bootstrap tests to examine the FHP for a panel of 12 EU countries over the period 1960-2002. Their results show that the bootstrap panel stability tests allow for cointegration between savings and investment in the long run with at least one break. Their country specific FMOLS estimates of β range from 0.59 to 1.03. Christopoulos (2007) employed panel Dynamic Ordinary Least Squares (DOLS) to estimate β with a panel of 13 OECD countries. For the whole period 1885–1992, the estimate of β is equal to 0.48, suggesting that the degree of mobility is relatively high among these countries. However, high capital mobility cannot be accepted for the sub-periods 1921-1992 and 1950-1992 (both are pre-Maastricht periods) where

[1] A similar finding that inclusion of additional variables like openness etc., affects estimates of β is also found in a forthcoming paper by Herwartz and Xu (2009).

the estimated values of β ranged from 0.79 and 0.90, respectively. Grier et.al (2008) examined the relationship between savings and investment in the USA using the Bai and Perron (1998, 2003) techniques to test for structural breaks. Using data from 1947Q1-2007Q1, their results show that the saving rate is stationary with two structural breaks in its mean and the investment rate is stationary without a break. By comparing the number of breaks and the pattern of mean shifts, they conclude that the US saving and investment rates are not linked in the long run. Their VAR-GARCH model showed a positive relation between the savings and investment rate in the short run. However, this relation has weakened dramatically over time in terms of both magnitude and statistical significance.

While Coakley et.al (2001), Katsimi and Moutos (2007) and Giannone and Lenza (2004) and Katsimi and Moutos (2007) have raised doubts on the validity of the FHP, others found that β is well below unity but decreased to about 0.5 or 0.4 in the post Bretton Woods and Maastricht periods lend some support for the existence of FHP in a much weaker form. In our view it is unlikely that in a changing and less than perfectly competitive dynamic international economic environment, a complete validity or invalidity of the FHP holds in all sample periods. Consequently, we think that perhaps the findings in the latter set of the above works that β was higher, and even close to the original estimates of FH in the pre Bretton Woods and Maastricht periods than in the post sample periods of these agreements is a more realistic conclusion. Therefore, in this paper we also test for structural breaks around 1972 for the effects of Bretton Woods and around 1992 for the effects of Maastricht agreements. We report estimates of β for the entire sample period with alternative panel data estimation methods as well as the relevant subsample periods. A problem that has been ignored in the panel data estimates with the levels of the variables, based on both the time series and classical methods, seems to be the likely presence of serial correlation in the residuals. We shall tackle this issue in Section 4.

3. System *GMM* and Structural Breaks

Generalized Method of Moments (GMM) is a semiparametrically efficient estimation method. Since Hansen (1982) established its large sample properties, GMM has gained a great deal of attention in the field of economics and finance over the past two decades. Although popular in economics it has been much used in finance area also. The GMM estimation methodology starts from a set of over-identified population of moment conditions and seeks to find an estimator that minimizes a quadratic norm of the sample moment vector. The resulting estimation has been shown to be consistent and asymptotically normal under suitable conditions.

Nevertheless, the GMM first-difference estimator suffers from a significant shortcoming. Blundell and Bond (1998) have shown that when the explanatory variables are persistent over time, lagged levels of these variables are weak instruments for the regression equation expressed in first-differences. Blundell and Bond (1998) also show that the instruments used with the standard first-difference GMM estimator (i.e. the endogenous variables lagged two or more periods) become less informative in models where the variance of the fixed effects is particularly high relative to the variance of the transitory shocks. This is likely to lead to biased coefficients, and the problem is generally exacerbated in small samples. To avoid this bias, Blundell and Bond (1998) proposed a system-GMM (henceforth SGMM) estimator. This estimator basically combines in a system the first-differenced with the same equation expressed in levels. The instruments for the regression in differences are the same as those described above, while the instruments for the equation in levels are lagged differences of the corresponding variables.

The main virtue of the SGMM approach consists in the fact that unlike *WITHIN* or *BETWEEN* (first-differences) approaches, it does use the estimation in levels for estimation and this exploits not only the variation in data over time but also between the countries. It thus allows to preserve more information to identify the parameters of interest. Arellano and Bond (1995) show on the basis of Monte-Carlo simulation that, this additional information results in a substantial gain in the precision of the estimation. Moreover, they set out that a sufficient additional condition (compared to the *GMM* estimator) for the validity of the *SGMM* estimator is to assume that the correlations between unobserved fixed effects and the explanatory variables are constant over time. It is also noteworthy to emphasize that the additional assumptions for the *SGMM* estimator do not affect the assumption of pre-determinedness of the inputs. As a consequence, the *SGMM* allows to control for simultaneity of input and output decision in the same way as the *GMM* estimator does.

Therefore, systems *GMM* estimator, introduced by Arellano and Bond (1995) and Blundell and Bond (1998), combines the standard set of equations in first differences with suitably lagged levels as instruments with an additional set of equations in levels with suitable lagged first differences as instruments (Bond et al., 2001). Thus, the consistency of the GMM estimates depends on the validity of the instruments. The validity of instruments that give a set of over-identifying restrictions has been verified with the standard Hansen test, which confirms that in all cases our set of instruments is valid. Furthermore, the *DW(1)* and *DW(2)* tests, that check the hypothesis of absence of serial correlation, are also presented. The standard errors of coefficients are robust to heteroscedasticity.

The puzzling finding by Feldstein and Horioka (1980) is the invariance of the saving-investment nexus to policy regime alterations towards capital mobility. Although it lies at the centre of the debate, incorporating the regime change effects into the analyses is yet to be a common practice. Moreover, since the capital mobility is known to have increased as a consequence of a worldwide shift towards financial liberalization (see *e.g.,* Frankel, 1992) any investigation of the existence of this relationship should allow for breaks. This point has been taken into account by both Banerjee and Carrion-i-Silvestre (2004) and Di Iorio and Fachin (2007), who applied different panel cointegration tests allowing for breaks. In our present paper, we investigate later the existence of structural breaks around 1972 and 1992, with the Mancini-Grifolli and Pauwels (2006) structural break test, for the effects of the Bretton Woods and Maastricht agreements.

The regression that serves as the basis for test of structural break is as follows:

$$Y_{it} = \begin{cases} X'_{it}\beta_0 + U_{it} & t = 1,\ldots,T \\ X'_{it}\beta_{1t} + U_{it} & t = T+1,\ldots,T+m \end{cases} \qquad (2)$$

for individuals $i = 1,\ldots,n$ and where T is the supposed break date. The test hinges the next hypotheses: $H_0 : \beta_{1t} = \beta_0$ against $H_A : \beta_{1t} \neq \beta_0$. In order to build the test the authors consider more observations after the break date than regressors d , $(m \times n) \geq d$. Briefly, the test statistic is a positive definite quadratic form obtained from the transformed $(m \times n) \geq 1$ vector of residuals by the $(m \times n) \times (m \times n)$ covariance matrix, projected onto the column space of $(m \times n) \times d$ matrix of transformed post-instability regressors. As the authors argue, the equivalent of the generic test statistic in Andrews (2003) for panel data can be defined after considering an interval τ_r which goes from $[r, r+m-1]$ and where $r \in \{1, ..., T+1\}$, as:

$$S_r(\beta, \Sigma) = A_r(\beta, \Sigma)' V_r^{-1} A_r(\beta, \Sigma), \tag{3}$$

$$A_r(\beta, \Sigma) = X'_{\tau} \hat{\Sigma}_{T+m}^{-1} \hat{W}_{\tau r}, \tag{4}$$

$$V_r(\Sigma) = X'_{\tau} \hat{\Sigma}_{T+m}^{-1} X_{\tau r} \tag{5}$$

with $\hat{W}_{\tau_r} = (\mathrm{Y}_{\tau_r} - \mathrm{X}_{\tau_r} \beta)$ where $\hat{W}_{\tau_r}$ is the $(m \times n) \times 1$ residual vector of observations starting at r, with $\beta = \hat{\beta}_{T+m}$ defined to be the coefficient vector estimated over the $T+m$. The variance-covariance matrix, $\hat{\Sigma}_{T+m}$, is given by:

$$\hat{\Sigma}_{T+m} = (T+1)^{-1} \sum_{r=1}^{T+1} \left(\hat{U}_{\tau} \hat{U}'_{\tau} \right) \tag{6}$$

where the $(m \times n) \times 1$ residual vector, $\hat{U}_{\tau_r}$, is defined as $\hat{U}_{\tau_r} = (\mathrm{Y}_{\tau_r} - \mathrm{X}_{\tau_r} \hat{\beta}_{T+m})$. It is noteworthy that this covariance matrix corrects for serially correlated errors, heteroscedasticity and potential cross-sectional correlation.

Hence, the test statistic for the post-break residuals is defined as:

$$S = S_{T+1}\left(\hat{\beta}_{T+m}, \hat{\Sigma}_{T+m} \right) \tag{7}$$

Accordingly, the critical values, S_r, are found by empirically, generating a distribution function for the statistic under the null of stability. As before, if $(m \times n) \geq d$ the $T-m+1$ different S_r values are defined as:

$$S_r = S_r\left(\hat{\beta}_{2,(r),}\, \hat{\Sigma}_{T+m}\right) \qquad (8)$$

where $\hat{\beta}_{2,(r)}$ is the estimate of β over $t = 1,...,T$ observations but excluding $\frac{m}{2}$ observations. The optimization of the power and size is the reason behind such exclusion, compared with the exclusion of only m observations or no observations at all.

However, the variance-covariance matrix , $\hat{\Sigma}_{T+m}$, as defined above will not be invertible in most cases, as it will in general not be of full rank, and thus for its adaptation to the panel data requires certain restrictions on the $(m\times n)\times(m\times n)$ covariance matrix to make it invertible. Therefore, the covariance matrix is redefined assuming sectional interdependence although continue to allow for serial correlation and heteroskedasticity. The redefined matrix has the following expression:

$$\tilde{\Sigma}_{T+m} = (T+1)^{-1}\sum_{r=1}^{T+1}\left(\hat{U}_{\tau_r}\hat{U}'_{\tau_r}\right) = \hat{\Sigma}_{T+m} \qquad (9)$$

except that $E\left[U_{i,\tau_r}\, U'_{j,\tau_r} \mid \mathrm{X}_{ij}\right] = 0$, for $i \neq j$ with $i, j = 1,...,n$ and U_{i,τ_r} is an $m\times 1$ vector made up of the elements in $U_{\tau r}$ corresponding to individual i. The resulting covariance matrix $\tilde{\Sigma}_{T+m}$ is block diagonal. Each block corresponds to an individual in the panel, and it is thus of dimension $(m\times m)$. Since the inverse of a block diagonal matrix is the inverse of each of its blocks, the condition for invertibility is satisfied[2].

It is worth noting that the aforementioned procedure for structural break testing offers three main practical and technical advantages over others. First, it does not make any distributional assumptions as it estimates empirically the distribution of the test statistic using an empirical subsampling methodology. Second, the power of the test remains high even when there are very

[2] See Mancini-Griffoli and Pauwels (2006) for detailed computations of alternative conditions for the inversion of the covariance matrix.

few observations after the break date. Third, the test requires very few regularity conditions. It remains asymptotically valid despite non-normal, heteroscedastic and/or autocorrelated errors, and non strictly exogenous regressors. We wish to highlight that among other tests, an important advantage of this one is that it does not require normal *iid* errors and strictly exogenous regressors, while the F-type tests do.

4. Empirical Results

Our sample includes 13 OECD countries for which data are available for 1960-2007 without any gaps. These countries are Australia, Belgium, Denmark, Finland, France, Great Britain, Germany, Greece, Ireland, Italy, Spain, Sweden and the USA. We report first estimates of equation (1) for the whole sample period with the standard panel data estimates viz., pure cross section or *TOTAL* estimates, 2 fixed effects models viz., *BETWEEN* and *WITHIN* and the random effects model *REM*. Second, we present the single equation estimate with *GMM* in which the first differences of the variables are used. This is the traditional approach with *GMM*. Finally, we shall use the systems *GMM* approach (*SGMM*) of Blundell and Bond (1998) in which the specifications in the first differences and levels of the variables are estimated simultaneously. Estimates with these alternative methods are given in Table 2. Two sets of subsample estimates with *SGMM* and *REM* only are reported to conserve space in Table 3.

Estimates with the country specific time series data and OLS for the whole sample period showed that there are some differences in the estimates of β between these 13 countries. It is highest at 0.885 for Italy and lowest at 0.266 for the USA and Belgium. For Ireland it is slightly higher at 0.328. For the rest of the countries, with the exception of France, the estimates are around 0.5. For France it is 0.711. These are not shown in Table 1 to conserve space. We only report estimates with panel data methods in Table 1. These range from 0.830 in column 2 with the fixed effects *BETWEEN* method to 0.461 in column 5 with the conventional single equation based *GMM* with the first differences of the variables. The rest of the estimates vary from 0.5 to 0.6. The *SBIC* selected the estimates with the *REM* in column 4 as the best among these 4 traditional panel data estimates. For reasons explained in the previous section the *SGMM* estimates in column 6 are to be preferred to single equation based *GMM* estimates in column 5.

The *SGMM* estimate of β at 0.570 is our preferred estimate for the whole sample period. On the basis of these results we may conclude that the *FHP* exists in a weaker form and as Sinha and Sinha (2004) have correctly observed the mother of all puzzles does not seem to go away. This conclusion is similar to the conclusions in the more recent studies by Fouquau et. al., (2009), Katsimi and Moutos (2007), Di Iorio and Fachin (2007) and Christopoulos (2007) where the estimates of β are about the same as ours.

Some of these authors have also estimated β for various subsamples although the selection of the subsample periods do not coincide with those believed to have affected this coefficient. These are the Bretton Woods and the Maastricht agreements of 1972 and 1992 respectively. Only Katsimi and Moutos's estimates for the subperiods 1996-2000 and 1997-2002 might have captured some effects of the Maastricht agreements. Their estimates of β for these subsample periods are much less at about 0.372 and 0.261, respectively, compared to their estimate of 0.572 for the total sample period of 1986-2002. Therefore, their results seem to suggest that the Maastricht agreement has significantly improved capital mobility.

Table 2

Conventional and *SGMM* Panel Estimates 1960-2007

$$ITY_{it} = \alpha_i + \beta_i STY_{it} + \varepsilon_{it}$$

	1	2	3	4	5	6
	TOTAL	*BETWEEN*	*WITHIN*	*REM*	*GMM*	*SGMM*
β_0	0.093 (0.00)	0.041 (0.32)	--	0.113 (0.00)	--	-0.022 (0.85)
β_1	0.592 (0.00)	0.830 (0.00)	0.493 (0.00)	0.501 (0.00)	0.461 (0.03)	0.570 (0.00)
SER	0.027	0.014	0.024	0.028	0.012	0.013
$\overline{R}^2$	0.406	0.624	0.548	0.406	0.120	1st differences equation 0.110 Levels equation 0.769
DW	0.200 (0.00)	--	0.249 (0.00)	0.184 (0.00)	2.456 (1.00)	1st differences equation 2.563 (1.00) Levels equation 0.031 (0.00)
ρ_1	--	--	--	0.251 (0.00)	0.407 (0.01)	0.414 (0.00)

In Table 3 the results for structural breaks with the Mancini-Grifolli and Pauwels (2006) tests are reported in the last row. We have tried with various break dates for the effects of both the Bretton Woods and the Maastricht agreements of 1972 and 1992 respectively. It is unlikely that these two agreements had instantaneous effects from 1972 and 1992 respectively. Therefore a reasonable lag of 3 years is assumed for their effects. Our selected subsample periods, therefore, are 1960-1974 and 1975-2007 for the Bretton Woods effect and 1960-1994 and 1995-2007 to capture the Maastricht effect. The results with the Mancini-Grifolli and Pauwels (2006) tests indicated that there has been a break in 1975 due to perhaps the Bretton Woods agreement. The alternative break dates that have been tried are 1974 and 1976 and the test results are similar. The computed test statistic for a break in 1975 is S=22.85 and exceeds the 1% critical value of S_r (1%)=21.67. Therefore, the null of no break in 1975 is rejected. Another set of break dates that have been tried are after 1992 for the effects of the Maastricht agreement. These dates are 1994, 1995 and 1996. In none of these dates there is a structural break. We report the test results for a break 1995 and the results for 1994 and 1996 are similar. The computed test statistic for a break in 1995 is S=13.17 which is less than the 1% critical value of S_r(1%)=55.08. Therefore, the null that there was no break in 1995 cannot be rejected. In addition to applying this test to *SGMM* estimates, we have also applied it to the *REM* estimates. The computed test statistic for a break in 1995 is S=19.82 and the critical value for 1% level is S_r(1%)=139.31. Therefore, the null that there was no break in 1995 cannot be rejected.

Estimate in column 1 of Table 3 are for the pre-Bretton Woods period where β at 0.963 is almost unity. The Wald test statistic for the null that $\beta = 1$, with the p-ratio in the brackets, is 0.037 (0.873) and the null cannot be rejected. It may be concluded that in the pre-Bretton Woods period there was almost zero or very little international capital mobility and the sources for investment were savings from the domestic sectors. Estimate of β for the post-Bretton Woods period in column 2 is dramatically less at 0.538. The computed Wald test statistic with the null that it equals the estimate in column 1 is 6.267 (0.012) and the null is rejected. Therefore, it can be concluded that the Bretton Woods agreement has significantly increased international capital mobility between the OECD countries in our panel.

Table 3

Systems *GMM* and *REM* Panel Data Estimates

1960-1974 & 1975-2007; 1960-1994 & 1995-2007

$$ITY_{it} = \alpha_i + \beta_i STY_{it} + \varepsilon_{it}$$

	1	2	3	4	5	6
	SGMM ESTIMATES				*REM ESTIMATES*#	
	1960-1974 Pre-Bretton Woods	1975-2007 Post-Bretton Woods	1960-1994 Pre-Maastricht	1995-2007 Post-Maastricht	1960-1994 Pre-Maastricht	1995-2007 Post-Maastricht
β_0	-0.011 (0.89)	0.032 (0.57)	0.030 (0.48)	0.113 (0.00)	0.098 (0.00)	0.117 (0.00)
β_1	0.963 (0.00)	0.538 (0.00)	0.528 (0.00)	0.289 (0.29)	0.590 (0.00)	0.414 (0.00)
$\overline{R^2}$	1st differences equation 0.230 Levels equation 0.833	1st differences equation 0.052 Levels equation 0.863	1st differences equation 0.083 Levels equation 0.820	1st differences equation 0.135 Levels equation 0.946	--	--
DW	1st differences equation 2.436 (1.00) Levels equation 1.212 (0.00)	1st differences equation 2.840 (1.00) Levels equation 0.489 (0.00)	1st differences equation 2.840 (1.00) Levels equation 0.270 (0.00)	1st differences equation 2.218 (1.00) Levels equation 1.170 (0.00)	--	--
ρ_1	0.795 (0.00)	0.764 (0.00)	0.646 (0.00)	0.995 [constrained]	0.296 (0.00)	0.746 (0.00)
SB	$S=22.85$; $S_r(1\%)=21.67$		$S=13.17$; $S_r(1\%)=55.08$		$S=19.82$; $S_r(1\%)=139.31$	

Notes: # These estimates are made with the maximum likelihood method unlike GLS in Table 2. TSP output does not compute the correlation coefficient and the DW statistic.

Estimate of β for the pre-Maastricht period in column 3 is almost the same in column 2. But we faced severe convergence problems while estimating for the post- Maastricht period in column 4. The estimated first order serial correlation coefficient ρ_1 was almost unity causing the convergence problem. Therefore, ρ_1 is fixed at 0.995 for the results in column 4 where β has decreased to 0.289 but insignificant at the 5% level. Changes to the starting date for this subsample, back and forth, produced very volatile results for β. Therefore, we reestimated this equation with the random effects model and the results are in columns 5 and 6. The estimate for β in column 5 for the pre-Maastricht period at 0.590 is slightly higher than the *SGMM* estimate of 0.528 of column 3. However, estimate of β for the post-Maastricht period at 0.414 in column 6 is significant and less than the estimate for the pre-Maastricht period. It was not possible to use the Wald test to say that β has significantly decreased in the post-Maastricht period because this option is not available for this test in the software we have used *(TSP*). Therefore, we have used the estimated standard errors to compute the 5% and 10% lower values of β for the pre-Maastricht period and upper value for the post-Maastricht period. While their 5% values overlapped slightly, their 10% values did not.[3] On the basis of this weak support we may say that Maastricht agreement at the most marginally increased international capital mobility, but this effect is not as large as the Bretton Woods effect. The structural break test also indicates that the Bretton Woods agreement has been more significant. Nevertheless, the FHP still survives but in a much weaker form after these two major international economic agreements.

5. Conclusions

In this paper we have attempted to fill a gap in the literature by applying a systems based *GMM* estimation method to test the validity of the mother of all puzzles namely the Feldstein-Horioka puzzle (FHP). We have also used a recently developed structural break test to understand the effects of two important international agreements viz., the Bretton Woods and the Maastricht agreements on international capital mobility. Our results showed that while the FHP is valid in

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[3] The computed 5% pre and post values, respectively, are 0.532 and 0.554 and they overlap only marginally. The 19% pre and post values are 0.541 and 0.529. Since the latter is less than the former it may be said that β has decreased marginally in the post- Maastricht period.

the pre-Bretton Woods period and international capital mobility was negligible, there has been a significant improvement in international capital mobility between the OECD countries in our sample in the post-Bretton Woods period. The effects of the Maastricht agreement on international capital mobility seem to be modest and far less than the Bretton Woods agreement. This distinction between the effects of these two important agreements somehow does not seem to have been made in the existing voluminous empirical literature on FHP. However, as noted at the end of the previous section this mother of all puzzles does not vanish and still exists in considerably weaker form. How to further improve international capital mobility is a sixty four dollar question and needs further investigation by the interested researchers. In light of the findings by Fouquau et.al (2009) that the degree of openness, the size of the country and the ratio of current account balance to GDP have significant effects on the estimates of β, it may be difficult to further improve international capital mobility between the OECD countries in our panel because these countries already are highly open economies with stable ratios of current account balances to GDP over longer periods. An alternative to get some insights into policies needed to improve capital mobility is country specific time series studies to highlight country specific rigidities against capital mobility.

However, our study and conclusions have limitations. Firstly, serial correlation seems to be a problem in all types of estimation methods to which not much attention has been given in the existing literature. In our paper we have only allowed for first order serial correlation and neglected higher order serial correlation. Secondly, the Mancini-Griffoli and Pauwels structural break tests assume prior knowledge of the break dates. Although throwing away such prior knowledge in favour of some endogenous structural break tests is methodologically controversial, the majority of researchers seem to prefer the latter. We are not aware of any endogenous structural break tests for the systems based *GMM* panel data methods. Therefore, hopefully some theoretical econometricians may pay attention to this gap. In conclusion, we hope that our results will receive further scrutiny and extension by others working on this mother of all puzzles.

Data Appendix

ITY is gross domestic investment as a share of GDP. Data obtained from International Financial Statistics (IFS) 2007.

STY is gross domestic savings as a share of GDP. Data obtained from IFS 2007.

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