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A Time-series Approach to the Feldstein–Horioka Puzzle with Panel Data from the OECD Countries

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1. INTRODUCTION

THE high correlation between domestic savings and investment is well known as the Feldstein–Horioka puzzle (henceforth FHP). It started with Feldstein and Horioka (1980, henceforth FH) where they have shown with the cross-section data of 16 OECD countries for the period 1960–74, that investment and saving ratios are highly correlated. Therefore, they argued that domestic saving is the main source of funds for investment, which in turn implies, according to them, that international capital mobility is low. However, this implication as evidence against capital mobility was questioned by some authors. Jansen (1996, 1998), Coakley and Kulasi (1997) and Pelgrin and Schich (2004) interpret the close long-run relationship between the investment and saving ratios as a solvency condition that must be satisfied and not as evidence against international capital mobility. Byrne et al. (2009) have argued that this correlation could be due to common global shocks and therefore it may not be interpreted as evidence against capital mobility. Nevertheless, we take the view of many that the FHP is a simple and an indirect test on the extent to which capital is mobile across countries. If tested for structural breaks it can also give an indication about changes in capital mobility. Capital mobility, in its own right, is important because it has implications for single currency debates, tax policies on capital and saving, whether growth is constrained by

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domestic saving and for the crowding effects of fiscal deficits. On the other hand, if capital mobility is high, countries cannot pursue independent monetary policies. Because of these policy implications Obstfeld and Rogoff (2000) have called FHP the mother of all puzzles.

In the FH cross-section regressions of the ratio of investment to gross domestic product (GDP) (investment ratio) on the ratio of saving to GDP (saving ratio), the coefficient of the saving ratio, known as the saving retention coefficient (β), was almost unity. This puzzle, despite a number of empirical investigations with alternative datasets, specifications and estimation technique, still remains a puzzle. The vast empirical literature on FHP is comprehensively surveyed by Apergis and Tsoumas (2009). They conclude that the majority of the empirical studies do not support the original strong results of FH but found that this correlation still exists in a weaker form in that β seems to have decreased and is significantly less than unity.

The outline of this article is as follows. Section 2 briefly reviews a few relevant empirical works. In Section 3, empirical results for panel unit root and cointegration tests and estimates of the cointegrating equations with tests for structural breaks are presented. Section 4 concludes.

2. BRIEF OVERVIEW OF PANEL STUDIES ON FHP

To test the validity of FHP, many studies estimate the following equation or its variants:

$$ITY_{it} = \alpha_i + \beta_i STY_{it} + \varepsilon_{it}, \tag{1}$$

where ITY_{it} is the domestic investment share of GDP and STY_{it} is the domestic saving share of GDP, i and t are country and time subscripts and $\varepsilon_{it} \sim N(0, \sigma)$ for all i and t. The null hypothesis is that, under complete capital mobility β in equation (1) should be zero. FH interpret this coefficient, also known as the saving retention coefficient, as an indicator of the degree of international capital mobility. Their empirical findings show that β is very close to one (between 0.85 and 0.95), indicating low capital mobility in the sample OECD countries. A number of recent panel data studies on FHP have focused on the OECD countries; for instance, see Cadoret (2001), Coakley et al. (2001, 2004), Giannone and Lenza (2004), Pelgrin and Schich (2004), Christopoulos (2007), Di Iorio and Fachin (2007), Herwartz and Xu (2009) and Fouquau et al. (2009).[1] Other recent panel data studies such as Coakley

[1] For discussions on cross-section and time-series studies on FHP, see Apergis and Tsoumas (2009). For our purpose, we review only recent empirical studies that utilised panel data estimation methods.

et al. (1999), Bahmani-Oskooee and Chakrabarti (2005), Kim et al. (2005), Chakrabarti (2006) and Murthy (2007a) are based on developing countries. The results in these studies considerably differ with some supporting and some against the validity of FHP.

The majority of the studies on OECD countries show that FHP exists in a weaker form because the savings–investment correlation is low. A brief summary of their findings is as follows. Di Iorio and Fachin (2007) employed the panel bootstrap tests to examine FHP for a panel of 12 EU countries over the period 1960–2002. Their country-specific Fully Modified Ordinary Least Squares (FMOLS) estimates of β range from 0.59 to 1.03. Christopoulos (2007) used the panel Dynamic Ordinary Least Squares (DOLS) to examine the FHP for 13 OECD countries. The estimate of β is around 0.5 for his whole sample period (1885–1992). However, for subsample periods (pre-Maastricht periods, i.e. 1921–92 and 1950–92) the estimated values of β ranged from 0.79 and 0.90, respectively. Giannone and Lenza (2004) utilised the Factor Augmented Panel Regression technique to examine the FHP for 24 OECD countries for the period 1970–99. This approach allows for heterogeneous response of savings and investment to global shocks. In the subsample period 1990–99, the relaxation of the homogeneity assumption reduced the estimate of β to 0.18. Coakley et al. (2001) utilised the time-series panel data techniques to examine the FHP for 12 OECD countries for the period 1980Q1–2000Q4. They obtain an estimate of β at around 0.32. Their findings support the integration of international financial markets in OECD countries. Coakley et al. (1999, 2004) also make similar findings on OECD and developing countries. Cadoret (2001) examined the FHP for 19 OECD countries for the period 1970–98 and found that β varies widely in different time spans.

Recently, Fouquau et al. (2009) have used the Panel Smooth Threshold Regression Model to test the validity of FHP for a panel of 24 OECD countries for the period 1960–2000. They included additional variables into the simple relationship between *ITY* and *STY* in equation (1) such as trade openness, the size of the country and the ratio of current account balance to GDP. Their estimates of β range from 0.5 to 0.7 and similar to the estimates of Herwartz and Xu (2009) for OECD countries. Pelgrin and Schich (2004) have used an error correction adjustment process and estimated dynamic fixed effects, mean group and pooled mean group equations for 20 OECD countries for the period 1960–99. They found that the error correction coefficient is negative and significantly different from zero. Pelgrin and Schich interpret the error correction coefficient as an indicator of capital mobility because a faster adjustment to equilibrium implies that the gap between *ITY* and *STY* is closed through international capital mobility. However, it is also possible that households and firms within a country respond by increasing the saving rate.

Further, numerous empirical studies have also analysed FHP for developing countries; however, only a few studies used panel data estimation methods. Bahmani-Oskooee and Chakrabarti (2005) have utilised the Pedroni's panel FMOLS technique to examine the savings–investment relationship for 106 countries. Their estimates of β are between 0.5 and 0.7. They found that β is significantly higher for the group of high-income countries than it is for the group of low-income countries. β is also higher for the group of closed economies than it is for the group of open economies. Similar findings are also made by Chakrabarti (2006) with a sample of 126 countries. Murthy (2007a) used the maximum likelihood panel cointegration techniques to examine the validity of FHP for 14 Latin American and four Caribbean countries over the period 1960–2002. His findings imply that correlation between savings and investment is very weak and the FHP is not valid in these countries. Kim et al. (2005) have estimated with time-series panel data methods β for 11 Asian countries for the period 1960–98. For the period 1960–79, they found that estimates of β are 0.58 and 0.76, respectively, with the FMOLS and DOLS methods. However, for the period 1980–98 estimates of β have decreased to 0.39 in FMOLS and to 0.42 in DOLS, implying that capital mobility has increased in the Asian countries. By contrast, Giannone and Lenza (2004) and Murthy (2007a) have found that there is no evidence to support for the validity of the FHP, the aforesaid studies and others have found that β is well below unity and provide some support for the existence of FHP in a weaker form.

However, in all these studies there were no formal tests for structural breaks in the relationship between saving and investment. Given that some major international agreements have been negotiated and accepted to increase globalisation to increase trade and capital mobility, it is likely that structural changes might have taken place in the relationship between investment and saving. In this article, we investigate this aspect of the FHP.

3. EMPIRICAL RESULTS

a. Unit Roots and Cointegration

Our sample comprises 13 OECD countries for which annual data are available from 1960 to 2007. These are Australia, Belgium, Denmark, Finland, France, Great Britain, Germany, Greece, Ireland, Italy, Spain, Sweden and the USA. The data on *ITY* (gross domestic investment as a share of GDP) and *STY* (gross domestic savings as a share of GDP) is obtained from *International Financial Statistics* (IFS) 2007.

We started through testing for the presence of unit roots in the two variables, namely *ITY* and *STY*, using the panel unit root tests of Breitung (2000), Levin

TABLE 1
Panel Unit Root Tests 1960–2007

Series	*LLC*	*Breitung*	*IPS*	*ADF*	*PP*	*Hadri*
ITY	−0.442	−3.209	−2.159	40.112	22.873	4.119
	(0.33)	(0.00)*	(0.02)*	(0.04)*	(0.64)	(0.00)*
STY	−1.217	−2.112	−1.144	33.337	25.187	6.853
	(0.11)	(0.02)*	(0.13)	(0.15)	(0.51)	(0.00)*
Δ*ITY*	−16.576	−10.169	−13.140	191.601	160.003	2.789
	(0.00)*	(0.00)*	(0.00)*	(0.00)*	(0.00)*	(0.00)*
Δ*STY*	−21.043	−11.796	−18.274	283.37	284.95	2.838
	(0.00)*	(0.00)*	(0.00)*	(0.00)*	(0.00)*	(0.00)*

Notes:
Probability values are reported in the parentheses. In LLC, Breitung, IPS, ADF and PP tests, the null is that the variable is non-stationary. However, in the Hadri test, the null is that the variable is stationary. For a discussion of these tests, see Baltagi (2005) and Pesaran and Breitung (2005).
*The rejection of the null at the 5% level.

et al. (2002, LLC), Im et al. (2003, IPS), ADF Fisher chi-square (ADF), PP Fisher chi-square (PP) and Hadri (2000). The panel unit root test results are given in Table 1.

These tests provide fairly mixed results for *ITY*. The LLC and PP tests in which the null is that the variable is non-stationary is not rejected at the 5 per cent level. However, the IPS and ADF tests in which the null is the same accept the null at only the 1 per cent level. In the Hadri test, the null is that the variable is stationary and it is rejected at the 5 per cent level. For *STY*, all the tests show that it is a non-stationary variable at the 5 per cent level, except for Breitung at the 1 per cent level. Alternatively, with the exception of the Hadri test, all other tests show that the first differences of *ITY* and *STY* are stationary. Therefore, it is reasonable to conclude that these variables are by and large I(1) in their levels.

The results of the panel cointegration tests and estimates of the panel cointegrating equations are reported in Table 2.[2] In the equations with common time trends, the majority of the cointegration tests, five of seven, show that there is cointegration between *ITY* and *STY* at the 5 per cent level. Only the panel v and group σ test statistics are insignificant at the 5 per cent level. The cointegration test results without common time trends are the other way, i.e. out of these seven tests only two (the panel v and group ADF test statistics) reject the null of no cointegration. However, it is well known that the two *ADF* tests have more power against the null and both reject the null of no cointegration in the model with common time trends, but in the model without the common

[2] The estimates of the individual country cointegrating parameters are relegated to the Appendix (see Table A1). For a comprehensive discussion on panel unit root tests and Pedroni's panel cointegration method, see Murthy (2007a, 2007b).

TABLE 2
Panel Cointegration Tests and Coefficients 1960–2007

Test Statistic/Savings Estimate (β)	*With Common Time Dummies*	*Without Time Dummies*
Panel ν-statistic	1.375	2.587*
Panel σ-statistic	−1.979**	−1.389
Panel $\rho\rho$-statistic	−2.751*	−1.248
Panel ADF-statistic	−3.438*	−1.479
Group σ-statistic	−0.809	−1.010
Group $\rho\rho$-statistic	−2.627*	−1.147
Group ADF-statistic	−4.512*	−2.049*
β	0.304	0.571
	(6.83)*	(13.90)*

Notes:
The test statistics are distributed as $N(0, 1)$. FMOLS estimates of β are reported where *ITY* is the dependent variable. The t-ratios are in parentheses. For a discussion of Pedroni panel cointegration tests, see Murthy (2007a, 2007b).
Significance at the *5% and **10% levels.

time trends the null is rejected by only one ADF test. Nonetheless, we can infer that *ITY* and *STY* are cointegrated and thus we estimate the cointegrating equation based on these two alternative models.

The estimates of β are around 0.3 and 0.6 in both types of models, respectively. This crucial savings retention coefficient is significant at the 5 per cent level. The country-specific estimates of β vary widely and this is not uncommon in the panel data studies.

b. Effects of Bretton Woods and Maastricht Agreements

We shall examine the effects of two important agreements which could have contributed to increased capital mobility, viz. the Bretton Woods and Maastricht Agreements. The Bretton Woods system of monetary management established the rules for financial relationships among the world's major industrial countries. This agreement started after World War II in 1945 and ended in 1972. Particularly this agreement established the pegging of currencies and the International Monetary Fund (IMF) in the hope of stabilising the global economic environment. The Maastricht Treaty began from 1992 between the members of the European Community. This agreement created the European Union and led to the creation of the euro. No doubt the effects of these two agreements on the savings–investment relationship are difficult to estimate. However, our sample period covers 13 years towards the end of the Bretton Woods agreement and the first 16 years of the Maastricht agreement. The former agreement can be said to have created exchange certainty and the latter, by integrating the political and economic union between the European Union countries, may have improved investor confidence. Both effects are likely to improve capital mobility, but this interpretation should be considered cautiously because alternative explanations for the decline in the saving–investment correlation are also possible. One such alternative explanation is Byrne et al. (2009). They have decomposed the correlation between the saving–investment relationship into two types, viz. correlation between country-specific factors and global factors. Their cointegration tests showed that the null of no cointegration could not be rejected between the country-specific factors but rejected for global factors. Since both variables in the saving–investment relationship are domestic variables, Byrne et al. conclude that capital mobility has increased. If significant correlations are found between saving and investment, they are due to common global factors and not due to decreased capital mobility.

For simplicity, we divided our sample into subsample periods to capture the effects of the Bretton Woods and Maastricht agreements. It is improbable that these two agreements had an instantaneous impact on capital mobility from 1972 and 1992, respectively. Hence, we assume that a lag of three years is reasonable for their effects. Consequently, we select subsample periods as 1960–74 (pre-Bretton Woods), 1975–2007 (post-Bretton Woods), 1960–94 (pre-Maastricht) and 1995–2007 (post-Maastricht). Prior to further discussion, it would be useful to take an overview of what is expected from these subsample estimates. Most importantly, we are investigating some evidence on whether the Bretton Woods and Maastricht agreements had any significant effects on the validity of FHP and capital mobility. If they have been effective, it is to be expected that the value of β will decline in the second set of subsamples to show an increase in the capital mobility.

The results of the cointegration tests and estimates of the cointegrating equations of the subsample periods are reported in Table 3. In the two sets of subsamples, the null of no cointegration is rejected by the more powerful ADF test statistics at the 10 per cent level, except for the model without common time trends in the post-Maastricht period. One or more of the other cointegration tests also confirm cointegration between *ITY* and *STY* at the 5 per cent level. The only exception is this model in the post-Maastricht period (1995–2007) where all cointegration tests do not reject the null of no cointegration. In light of the above observations, we assert that there is no strong evidence that there is no cointegration in the two sets of subsample periods, except in the model without common trends in the post-Maastricht period.

The pre-Bretton Woods period highlights that the estimate of β is 0.467 and 0.742, respectively, in the models with and without common time trends. In both models, the estimate of β has decreased to 0.266 and 0.486, respectively, in the post-Bretton Woods period. Similar results are also found between the

TABLE 3
Panel Cointegration Tests and Coefficients: Subsamples

Test Statistic/ Savings Estimate (β)	*Pre-Bretton Woods 1960–74*		*Post-Bretton Woods 1975–2007*		*Pre-Maastricht 1960–94*		*Post-Maastricht 1995–2007*	
	With Common Time Dummies	*Without Time Dummies*	*With Common Time Dummies*	*Without Time Dummies*	*With Common Time Dummies*	*Without Time Dummies*	*With Common Time Dummies*	*Without Time Dummies*
Panel v	0.873	1.574	0.293	1.571	1.540	4.562*	−0.117	0.640
Panel σ	0.857	−0.580	−0.365	−1.358	−1.573	−2.457*	1.698**	0.221
Panel $\rho\rho$	−0.482	−1.117	−1.440	−1.730**	−2.276*	−1.930**	−0.349	−0.370
Panel ADF	−2.470*	−1.870**	−1.748**	−1.945**	−3.803*	−2.866*	−3.140*	−0.723
Group σ	2.255*	1.021	0.807	−0.004	−0.209	−1.583	2.855*	1.628
Group $\rho\rho$	0.236	−0.379	−0.804	−1.051	−1.579	−1.652**	−0.099	0.460
Group ADF	−4.403*	−1.855**	−2.281*	−1.978*	−4.167*	−3.372*	−3.547*	−0.147
β	0.467	0.742	0.266	0.486	0.443	0.652	0.248	0.115
	(12.82)*	(13.73)*	(6.08)*	(8.70)*	(9.34)*	(17.48)*	(7.01)*	(5.52)*

Notes:
t-Ratios are in parentheses. For the post-Maastricht period, without trend, all the cointegration tests did not reject the null of no cointegration at the 10% level; however, the group σ-statistic is close to 10% significance level. FMOLS estimates of β are reported where *ITY* is the dependent variable. For a discussion of Pedroni panel cointegration tests, see Murthy (2007a, 2007b).
Significance at the *5% and **10% levels.

pre- and post-Maastricht periods. The estimate of β has decreased from 0.443 to 0.248 in the model with time trends and from 0.652 to 0.115 in the model without trends. The country-specific estimates of β based on the subsample periods are not reported but are available from the authors upon request. These results show that for majority of the OECD countries, the estimates of β have slightly declined due to the Bretton Woods and Maastricht agreements, thus implying that international mobility of capital has marginally increased in these countries.

We have also tested for structural breaks using the Westerlund (2006) method and the results are reported in Table 4. This helps to verify if our choice of the above dates is reasonable.

We tested for two breaks and found that in Belgium, Finland, Germany and Greece there is only one break, but in the other countries there are two breaks. The results in Table 4 indicate that from the late 1960s to the early 1970s there have been structural breaks in Denmark (1966), Australia (1972), Great Britain (1970) and Italy (1970). In the other countries, the break occurred later in the late 1970s and early 1980s. These countries are Belgium (1981), France (1980), Greece (1983), Ireland (1981), Spain (1983) and the USA (1977). In Germany and Sweden, the break seems to have taken place in the late 1980s. There is thus a mixed result that the Bretton Woods agreement had a uniform effect on all the OECD countries to improve capital mobility. This prolonged period for structural adjustments may be due to the differences in the response by these countries to the economic uncertainties of the early 1970s. During this period, the Bretton Woods fixed exchange rate system collapsed and was replaced with different managed exchange rate systems. There were high inflation and severe energy crises, which in turn encouraged more conservative budgetary and monetary policies as well as some market liberalisation policies. Therefore, an improvement in the international capital mobility seems to have taken place over a longer time span and at different times in different countries.

TABLE 4
Westerlund Tests for Structural Breaks 1960–2007

Country	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*	*(7)*	*(8)*	*(9)*	*(10)*	*(11)*	*(12)*	*(13)*
Number of breaks	2	1	2	1	2	2	1	1	2	2	2	2	2
Break dates	1972	1981	1966	1992	1980	1970	1989	1983	1981	1972	1983	1988	1977
	1990		1989		1996	1990			1988	1992	1992	1995	1990

Notes:
Countries are numbered from (1) to (13), for example (1) Australia, (2) Belgium, (3) Denmark, (4) Finland, (5) France, (6) Great Britain, (7) Germany, (8) Greece, (9) Ireland, (10) Italy, (11) Spain, (12) Sweden and (13) USA.

By contrast, the dates for the second break are more uniform and around the late 1980s and the early 1990s. A second structural break occurred in nine of the 13 OECD countries, and these are Australia (1990), Denmark (1989), France (1996), Great Britain (1990), Ireland (1988), Italy (1992), Spain (1992), Sweden (1995) and the USA (1990). We have also tested for a single-structural break, but these results are not shown in Table 4 to conserve space. The results showed that there was a break during the late 1980s and early 1990s except in Greece and Ireland. There is no evidence that there was a break in the early 1970s. It may be recalled that the estimates of β in both the post-Bretton Woods and post-Maastricht agreements are about 50 per cent lower than in the pre-agreement periods. Based on the Westerlund tests, we may conclude that the Maastricht agreement seems to have had a more uniform and widespread effect on improving capital mobility in the major OECD countries. The lower estimate for β in the post-Bretton Woods period may be due to the inclusion of the period for the post-Maastricht period in this sample.

4. CONCLUSION

In this article we have used the time-series-based panel data methods and data from 13 OECD countries to test the validity of the mother of all puzzles, viz. the FHP. FHP has stimulated a large number of empirical works because of its important implications. It has directly or indirectly implied that international capital mobility was very low even among the advanced capitalist OECD countries. While this finding of FH's seminal contribution might be valid for their sample period of the 1960s and up to the collapse of the Bretton Woods agreement in the early 1970s, subsequently the turmoil caused by the collapsed fixed exchange rate system and the economic uncertainties of the 1970s led to the implementation of liberalisation policies and reforms, which seems to have improved the international capital mobility. However, the Maastricht agreement of the early 1990s has significantly improved international capital mobility. The saving retention coefficient is halved and less than 0.25 now.

However, our study and conclusions have some limitations. First, the break dates in the Westerlund tests are somewhat sensitive to the selected method of estimation and the number of breaks selected. Second, due to data limitations we have included only 13 OECD countries in our sample. Third, the scope of the software used for the Pedroni estimation method is limited in that it is not possible to use the Wald-type chi-squared tests to test restrictions on the coefficients. Nevertheless, our conclusion that there have been significant structural breaks in the FH equation and international capital mobility has improved in the post-Bretton Woods and Maastricht periods seems to be robust and valid.

APPENDIX

TABLE A1
Pedroni's Country-specific Estimates 1960–2007

Country	*With Common Trends β (t-ratios)*	*Without Trend β (t-ratios)*
Australia	0.293	0.544
	(2.48)*	(8.19)*
Belgium	0.230	0.225
	(2.05)*	(1.00)
Denmark	0.201	0.469
	(1.27)	(2.37)*
Finland	0.141	0.625
	(0.47)	(1.96)**
France	0.410	0.769
	(4.38)*	(6.45)*
Great Britain	−0.052	0.490
	(0.26)	(3.95)*
Germany	0.791	0.765
	(3.50)*	(5.95)*
Greece	0.337	0.525
	(3.03)*	(5.82)*
Ireland	0.534	0.484
	(3.05)*	(2.11)*
Italy	0.379	1.025
	(2.35)*	(5.28)*
Spain	0.317	0.709
	(0.59)	(2.20)*
Sweden	0.236	0.566
	(1.13)	(2.10)*
USA	0.138	0.226
	(0.60)	(2.72)*

Notes:
t-Ratios are in parentheses.
Significance at the *5% and **10% levels.

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