The Feldstein-Horioka puzzle is not as bad as you think

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Abstract

A country's intertemporal budget constraint implies current account stationarity or that its saving and investment rates should cointegrate. However such behavior may not be observed in finite samples where the current account could be subject to persistent shocks. Accordingly, this paper reconsiders the Feldstein-Horioka puzzle in a nonstationary panel framework for a sample of 12 OECD economies 1980I-2000IV. The mean group procedure gives a slope coefficient estimate which is insignificantly different from zero at the usual 5% significance level. This supports long run capital mobility and the globalization of international financial markets despite persistence in the current account.

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1 Introduction

Feldstein and Horioka (1980) (FH hereafter) established that long run averages of national saving and domestic investment — both expressed as ratios to GDP — were highly correlated in a cross section regression for 16 OECD economies 1960-74.[1] They interpreted this high association as implying segmented capital markets or low capital mobility. In a world of unfettered capital mobility, national saving would flow to the countries offering the highest returns and domestic investment could be financed from global capital markets. This is the basis for their reasoning that high capital mobility should imply a low saving-investment association and vice versa. The FH puzzle — which Obstfeld and Rogoff (2000) identify as one of the six major puzzles in international macroeconomics — refers to the stylized empirical finding that subsequent estimates of this saving-investment association have remained stubbornly high despite ongoing financial market liberalization and globalization in recent decades.[2]

The puzzle continues to exercise the imagination of economists as exemplified by the stream of very recent contributions.[3] It does so despite claims that saving-investment correlations are not informative about capital mobility since both are subject to cyclical shocks.[4] The latter overlooks the fact that the Feldstein-Horioka debate mostly has focused on the low frequency and not the business cycle component of the data. The continuing interest in the puzzle raises the obvious question of whether the apparently high capital mobility of recent decades is a chimera or an elusive reality. This issue matters since capital mobility is critical both for the efficient allocation of capital to the most productive uses and locations and for consumption smoothing. It is also relevant for policy issues such as the European Union single currency

[1]Hereafter we follow the FH literature in referring to both variables as a proportion of GDP as saving and investment. Feldstein (1983) extended the FH sample up to 1980 and confirmed their results.

[2]See Coakley, Kulasi and Smith (1998) and Obstfeld (1994) for recent surveys.

[3]Blanchard and Giavazzi (2002), Coiteux and Olivier (2000), Corbin (2001), De Vita and Abbott (2002), Ho (2002), Hoffman (2001), Isaksson (2001), Jansen (2000), Kim (2001), Kraay and Ventura (2002), Mark, Ogaki and Sul (2003), Obstfeld and Rogoff (2000), Obstfeld and Taylor (2002), Ozmen and Parmaksiz (2003), Sachsida and Caetano (2000), Schmidt (2001), and Taylor (2002). Due to space considerations, this list is confined to papers appearing in the new millennium. We apologise to any author(s) whom we may have unwittingly omitted.

[4]See for example Baxter and Crucini (1993).

debate (Bayoumi, Sarno and Taylor, 1999), large current account deficits in the Euro area (Blanchard and Giavazzi, 2002), and the role of net overseas balances (Lane and Milesi–Feretti, 2001).

Since the current account is the difference betwen saving and investment, the FH puzzle can be reformulated as the question of why incremental saving is invested domestically rather than overseas, leaving the current account unchanged. Many recent FH studies involve tests of current account stationarity or of cointegration between saving and investment. Hence, they implicitly rule out any role for permanent innovations. This paper marks a departure from such approaches. Its first contribution is that it provides a new measure of the long run saving-investment association using a panel regression approach that can accommodate persistent shocks to the current account. Such shocks may be theoretically plausible in finite samples since the transversality condition implied by the long run budget constraint applies only in the limit. Moreover, in reality countries do seem to run persistent current account deficits or surpluses as the recent experiences of the US and Canada testify and the data routinely reject the predictions of the present value model. Nason and Rogers (2000) summarize the situation as follows: 'Thus, the literature has reached the point where the intertemporal approach, although rejected in its most basic form, is still viewed as "useful" overall.' (p.1). They use Bayesian Monte Carlo experiments to show that physical trading costs and shocks to fiscal policy and world real interest rates may lie behind these rejections.[5]

Our analysis seeks to accommodate such shocks by exploiting recent advances in the nonstationary panel literature by Phillips and Moon (1999, 2000) and Kao (1999). These studies provide asymptotic results which demonstrate that in panels it is possible to estimate consistently the long run average association between two nonstationary variables even in the absence of cointegration between them.[6] The FH puzzle provides a convenient application since, while saving and investment are persistent or nonstationary processes, the evidence on cointegration is weak or mixed as emphasized in most recent contributions (Coakley, Kulasi and Smith, 1996; Coiteux and Olivier, 2000; Ho, 2002).[7]

[5] Note that Obstfeld and Rogoff (2000) also identify trading costs as a theoretical explanation of the FH puzzle.

[6] Coakley, Fuertes and Smith (2001) provide small sample evidence of this result in a Monte Carlo investigation.

[7] Taylor (2002) is an exception since he establishes current account stationarity in all

The second contribution is that our approach permits a high degree of country heterogeneity. One of the original criticisms of the FH approach was that their cross section estimator ignores heterogeneity across groups such as country size (Murphy, 1984). Moreover it does not exploit the time series variation either since the variables are averaged over the sample time span. The importance of country heterogeneity has been stressed for the panel literature generally (Boyd and Smith, 2000) and for the FH puzzle specifically (Corbin, 2001; Taylor, 2002). This paper examines the long run saving-investment association in a panel of 12 OECD economies for the 1980I-2000IV period using the mean group (MG) estimator proposed by Pesaran and Smith (1995). This panel approach to estimating long run relationships offers two advantages. On one hand, it incorporates country heterogeneity by allowing for both country-specific intercepts and slopes in the FH regressions. On the other, it has the merit of being able to sidestep the conceptual problems of panel unit root and cointegration testing. For the former, these include the difficulties in formulating the appropriate hypotheses for such tests (Baltagi and Kao, 2000) and potentially misleading power properties when a just fraction of the series is stationary (Karlsson and Löthgren, 2000).

The remainder of the paper is organized as follows. Section 2 discusses the link between the FH and intertemporal current account approaches. Section 3 outlines the econometric framework and Section 4 analyses the results. A final section concludes.

2 Intertemporal current account approach

In recent years the intertemporal current account approach or present value model has become the dominant theoretical framework for analysing the FH puzzle. This framework applies basic ideas from dividend discount models to a country's net asset position. In the former the fair price of an equity is the expected present value of future cash flows or dividends.[8] By analogy, a country's equilibrium net asset position is the expected present value of future cash flows from net exports. This can be derived by starting from the

cases for his panel of 15 countries but uses a long span of more than a century of annual data.

[8] See Campbell, Lo and MacKinlay (1997).

national income identity for gross national product:[9]

$$Y \equiv C + I + G + NX + rB \tag{1}$$

where C is private consumption, I is domestic fixed investment, G is public consumption, NX is net exports or exports less imports, r is the world rate of interest and B is the country's net asset position. The link between saving and investment and the current account is given by:

$$Y - C - G - I = S - I = CA \equiv NX + rB \tag{2}$$

where $S \equiv Y - C - G$ is national saving. From these identities, one can obtain closed form solutions to the forward-looking difference equations for a country's net asset position by assuming zero growth or constant growth analogously to solving dividend discount models.

First assume zero output growth. Given the equality of the capital and current accounts $B_t - B_{t-1} = -KA_t = CA_t$, a country's net asset position at time t can be written in difference equation form as:

$$B_t = R_t B_{t-1} + NX_t \tag{3}$$

where $R_t = 1 + r_t$. Now assume the interest rate factor is a particular martingale process, $E(R_{t+j} \mid \Theta_{t-1}) = R > 1$ for $j \geq 0$, where Θ_{t-1} is the latest available information set. Equation (3) can iteratively be solved forward to give:

$$B_{t-1} = -\sum_{j=0}^{\infty} R^{-(j+1)} E(NX_t \mid \Theta_{t-1}) + \lim_{j \to \infty} R^{-(j+1)} E(B_{t+j} \mid \Theta_{t-1}) \tag{4}$$

where the first term states that a country's net asset position is determined by the expected present value of its future net exports. The long-run budget constraint (LRBC) hypothesis is generally formalized as the limit condition that the last term in (4) must equal zero. In the analogous theory for the fair price of an equity, this transversality condition rules out speculative bubbles.

Now assume that output grows at rate g_t. In this case, scaled variables are used in the difference equation:

$$b_t = \frac{1 + r_t}{1 + g_t} b_{t-1} + nx_t \tag{5}$$

[9] While this section draws on the equations in Taylor (2002), the finance analogy is our own.

where lower case variables are defined as shares of output. Assuming $\frac{1+r_t}{1+g_t}$ to be the martingale process, $E(\frac{1+r_{t+j}}{1+g_{t+j}} \mid \Theta_{t-1}) = \rho > 1$ for $j \geq 0$, the LRBC now rules out explosive behaviour in the scaled future net asset position by letting its present value tend to zero in the limit. Note that, in finite samples, the LRBC may not hold exactly in both the zero and positive growth scenarios.

The LRBC has a straightforward econometric implication. Since saving and investment behave as nonstationary $I(1)$ processes — on which there is consensus — they should form a $(1, -1)$ cointegrating vector or, equivalently, the current account is a stationary process. Many of the recent contributions to the FH debate adopt this tack and test for cointegration between saving and investment or whether shocks to the current account are transitory (Coakley and Kulasi 1997; Coakley, Hasan and Smith 1999; Coakley, Kulasi and Smith 1996; Grundlach and Sinn 1992; Jansen 1996, 1997; Jansen and Schulze 1996; Taylor 1996, 2002).[10] However, in the case of a stationary current account, the implied high long run association between saving and investment may or may not be informative about long run capital mobility.

Moreover, the LRBC assumed in the theoretical models of the current account is a limit or large T condition. Thus it does not preclude persistent behaviour in finite samples. This is illustrated by the prolonged current account imbalances of many countries in recent decades. There are also good theoretical reasons to expect persistent deviations from current account balance. For instance, Herbertsson and Zoega (1999) argue that national-income identities and the life-cycle theory of consumption together imply that the current account should be a function of demographics, specifically the age structure. Using a panel of 84 countries, they find empirical support for the hypothesis that a country with a high proportion of young and retired should have protracted current account deficits.

The upshot of the above discussion is that in finite samples the current account may be observationally equivalent to a nonstationary process. Hence it is important to cater for persistent shocks or near $I(1)$ innovations to the current account in measuring the saving-investment association. This is particularly important when the data span is limited and comprises just a few decades which is typical of FH studies. The study of Taylor (2002) is

[10] This approach also leads to a distinction between long run and short run measures of the saving-investment association (Coiteux and Olivier 2000; Jansen 1996; Kraay and Ventura 2002; Sarno and Taylor 1998; Taylor 2002). Since the original FH finding relates to the long run association, this remains the focus of this paper.

a notable exception in this respect since he analyses capital mobility over the course of some 120 years (1870-1990) for a panel of 15 countries in a vector error correction framework.[11] Unsurprisingly with such a long span, he finds that the current acount is stationary or that the LRBC holds in all 15 countries.

Nonetheless Taylor adds the rider: "This is not to say that, in some periods, countries were unable to run "unsustainable" current account deficits, which were occasionally disrupted by crisis, real adjustments, or defaults. Episodes in some countries during the 1890s, 1930s, or 1980s could fit this description..." (*Ibid.* p.6). We conjecture that the typical data spans employed to address the FH puzzle may be picking up elements of such episodes so that the current account may appear nonstationary. Below we outline a methodology to deal with this issue.

3 Panel estimation of long run relationships

Given the mixed evidence on the cointegration of saving and investment and on the present value model of the current account, it seems appropriate to allow for the possibility of non-cointegration at the outset. In a time series regression involving two $I(1)$ variables such as saving and investment, the absence of cointegration leads to the well-known statistical problem of spurious correlation. However, recent work by Phillips and Moon (1999) and Kao (1999) shows that panel datasets offer the prospect of overcoming the spurious regression problem of pure time series. More particularly, they demonstrate that in large N, large T panels one can obtain consistent estimates of a long-run average parameter even if there is no time-series cointegration at an individual level or, equivalently, when the error term as well as the variables are nonstationary. The intuition is that the averaging over i (countries in the FH case) lessens the noise in the relationship — the covariance between the $I(1)$ error and the $I(1)$ regresssor — that induces the nonsense regression problem and leads to a stronger overall signal than in the pure time series regression case.

The Phillips and Moon insights and related work can be illustrated in the

[11] For a fascinating historical overview of capital mobility in the context of globalization over the same period, see Obstfeld and Taylor (2002).

context of the the following panel regression model:

$$y_{it} = \alpha_i + \beta_i x_{it} + u_{it}, \ i = 1, ..., N, \ t = 1, ..., T \tag{6}$$

where y_{it} and x_{it} are both $I(1)$ and suppose that u_{it} is also $I(1)$ so that y_{it} and x_{it} are not cointegrated. The fixed effects (FE) or within estimator, which assumes homogeneous slopes $\beta_i = \beta$, is defined by

$$\hat{\beta}^{FE} = \frac{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{x}_{it} \tilde{y}_{it}}{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{x}_{it}^2} = \beta + \frac{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{x}_{it} u_{it}}{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{x}_{it}^2} \tag{7}$$

where $\tilde{x}_{it} = x_{it} - \bar{x}_i$ and $\bar{x}_i = T^{-1} \sum_{t=1}^{T} x_{it}$ and similarly for $\tilde{y}_{it}$. In a time series setup, the noise, $\sum_t \tilde{x}_{it} u_{it}$, swamps the signal, β, and hence the OLS estimator will not converge to the true β as T becomes large. However, this problem is alleviated in a panel context by averaging over i and so a consistent estimate of β can be obtained as $N \to \infty$ and $T \to \infty$. Hence, the cross section variation in the data may help to extract the signal in nonstationary panels.[12]

To establish the applicability of these asymptotic results, Coakley, Fuertes and Smith (2001) use Monte Carlo simulations to investigate the finite sample properties of three panel regression estimators in the context of $I(1)$ errors. These are two pooled estimators — FE and pooled OLS (POLS) — and the mean group (MG) estimator of Pesaran and Smith (1995).[13] Coakley et al. show that, in a static regression with $I(1)$ errors for panels of the typical dimensions used in applied work, the above panel estimators appear unbiased with dispersion that falls at rate $\sqrt{N}$ even when the error term is $I(1)$. Moreover, standard t-tests for the MG estimator based on the $N(0, 1)$ distribution have reasonably good size properties irrespective of $I(0)$ or $I(1)$ errors. By contrast, inference based on the FE and POLS estimators is likely to be misleading since the usual standard error formulae are incorrect (leading to severely oversized t-tests) both in the $I(1)$ error case and in the heterogeneous slope case with $I(0)$ errors.[14] These are precisely the cases which are relevant for the FH puzzle.

[12] One caveat is in order. In line with most panel data work this asymptotic theory rests on the assumption of uncorrelated disturbances across groups. Little is known about the joint effect of $I(1)$ errors and between-group dependence.

[13] Asymptotic results have not been established for the MG estimator. However, this Monte Carlo investigation shows that it has somewhat better small sample properties than the pooled estimators.

[14] Phillips and Moon (1999) derive correct standard errors for the FE and POLS estimates for a number of DGPs.

Recall that FH originally used a cross section (CS) or between estimator to measure the long run saving-investment association:

$$I_i = \alpha + \beta^{CS} S_i + u_i, \quad i = 1, ..., N. \tag{8}$$

where i is a country index and u_i is a random innovation. The data points are long period averages, $I_i = T^{-1} \sum_{t=1}^{T} I_{it}$ and $S_i = T^{-1} \sum_{t=1}^{T} S_{it}$, where I_{it} is domestic fixed investment and S_{it} is national saving both defined as a ratio of GDP. FH employed this estimator, which exploits only the cross-section variation in the sample, to rule out comovement caused by short run or common cyclical influences. As such, they interpreted their regression coefficient as a measure of the long run saving-investment association.

The long run focus of FH's seminal paper has generally been followed in the literature by a succession of cross section studies and, more recently, unit root and cointegration studies (Coakley et al. 1998). One critical aspect of the latter is that they build on the premise that this high association is tantamount to a stationary current account. However factors such as productivity and demographic shocks may well induce $I(1)$ behaviour in the error term of a regression of investment on saving. The latter could explain the poor results from studies of the cointegration of saving and investment and, equivalently, of the present value model of the current account. Hence, an important issue is how to measure the long run saving-investment coefficient if the errors appear to be nonstationary or observationally equivalent to $I(1)$ processes.

Pesaran and Smith (1995) noted that the problem of spurious correlation does not appear in a CS regression. They show that, under the strong restrictions of random parameters and strictly exogenous regressors, the CS estimator is still consistent even if the errors are $I(1)$. However, the finite sample properties of the estimator in this nonstationary context remain to be established. The CS approach also disregards heterogeneity across countries. More specifically, suppose that the true unknown DGP has heterogenous coefficients as in (6) then the CS regression residuals measure $v_{it} = (\alpha_i - \alpha) + (\beta_i - \beta)\bar{x}_i + u_{it}$. If the individual unobservable attributes captured by the latter are correlated with the explanatory variable $\bar{x}_i$, then this will render the CS estimator inconsistent. Hence, CS estimates need to be interpreted with caution.

Recent contributions (Krol 1996; Coiteux and Olivier 2000; Corbin 2001; Jansen 2000) employ fixed effects (FE) estimators to tackle the heterogeneity issue. These introduce some heterogeneity by allowing for country and/or time dummies but still impose a common slope coefficient. In the context of nonstationary variables the last restriction will lead to downward biased standard errors and hence severely oversized t-statistics. Futhermore, size problems will arise in the absence of cointegration irrespective of homogeneous or heterogeneous slopes. Extant saving-investment studies have not considered these potential biases in their FE-based inference. This is precisely the issue on which the Phillips and Moon (1999) and Coakley et al. (2001) contributions shed light.

This paper employs the Pesaran and Smith (1995) MG approach to reassess the long run saving-investment association in a panel framework which accommodates both permanent current account shocks and heterogeneity across countries. Accordingly, the following regression — allowing for country-specific intercepts and slope coefficients — is run separately for each country by OLS:[15]

$$I_{it} = \alpha_i + \beta_i S_{it} + u_{it}, \; i = 1, ..., N, t = 1, ..., T. \tag{9}$$

to obtain the individual slope estimates $\hat{\beta}_i$. The MG estimate and its standard error are calculated as follows:

$$\widehat{\beta}^{MG} = \bar{\beta} = \sum_{i=1} \hat{\beta}_i / N \tag{10}$$

$$se(\widehat{\beta}^{MG}) = \sigma(\hat{\beta}_i)/\sqrt{N} \tag{11}$$

where

$$\sigma(\hat{\beta}_i) = \sqrt{\sum_{i=1} (\hat{\beta}_i - \bar{\beta})^2/(N-1)} \tag{12}$$

The MG slope estimator provides a measure of the average long run saving-investment association.

4 Data and results

Quarterly observations on national saving, domestic investment and GDP for the period 1980I-2000IV are gathered for 12 OECD countries.[16] These com-

[15]The time series estimates of the slope coefficient reported in the FH literature are markedly heterogeneous (Coakley et al., 1998).

[16]The sample size was dictated by data availability at a quarterly frequency since 1980.

prise Australia, Canada, Finland, France, Italy, Japan, Netherlands, Norway, Spain, Switzerland, UK and US. This panel is quite heterogeneous ranging from the US and Japan on one hand to Finland and Norway on the other. Nonetheless, European economies dominate since they account for 8 of the total. However it is noteworthy that our sample excludes European countries such as Greece, Ireland, and Portugal whose anomalous current account behaviour is highlighted in Blanchard and Giavazzi (2002).

We first test the null hypothesis that saving and investment are integrated or unit root processes. Table 1 presents the ADF test results where the augmentation lag (k) is selected by a testing down procedure at the 10% level starting from $k_{\mathsf{max}} = 8$.

[Table 1 around here]

It shows that the unit root null cannot be rejected for virtually all the saving and investment series. These findings are in line with the existing evidence that saving and investment — although expressed as ratios of output and thus bounded — mimic the properties of $I(1)$ processes.

Next the time series properties of the innovation term in the individual group regressions is investigated. We test the null of nonstationarity or that the saving-investment relationship is affected by permanent shocks. The results from the augmented Engle-Granger statistic in Table 1 suggest that the innovations in (9) are observationally equivalent to $I(1)$ processes, notwithstanding its shortcomings in small samples. This motivates the use of the MG regression estimator which appears to have good size properties in the presence of either $I(0)$ and $I(1)$ errors.

Table 2 presents two panel estimates of the long run saving-investment association in the FH regression. First, it reports the long run CS estimate and associated t-statistic for a comparison with the FH literature. In the former, the data are averaged over the full sample period and a cross section regression is estimated for these long averages. Second, it presents the time series MG estimate and inference to cater for possibly $I(1)$ innovations.

[Table 2 around here]

The results are revealing. Our CS estimate of 0.68 conforms to the "two-thirds" rule in the FH literature for the proportion of a saving increment invested domestically (Coakley et al., 1998; Table 1). It is still larger than one might anticipate, in the light of ongoing capital market integration, given that it is based on data for recent decades (1980-2000). The CS scatterplot is illustrated in Figure 1.

[Figure 1 around here]

The plot suggests that the results are not driven by outliers.[17] Moreover, the hypothesis that the average long run association is zero ($\beta = 0$) is strongly rejected in line with the FH literature.[18] Thus these findings appear to support the FH puzzle of low long run capital mobility.

By contrast, the average long run $\hat{\beta}^{MG}$ estimate of 0.33 is less than half the CS estimate and the hypothesis of perfect capital mobility ($\beta = 0$) cannot be rejected at the usual 5 per cent significance level.[19] This finding is in agreement with the notion that capital is highly mobile in the long run which is precisely the opposite inference to that from the CS approach. To our knowledge, this is the first time an insignificant panel estimate of the long run saving-investment association has been reported in the FH literature.

Our findings decisively overturn the existing evidence and is in line with perceptions of high capital mobility in the closing decades of the twentieth century. They thus exemplify Taylor's (2002) view of recent decades marking a return to the degree of capital mobility witnessed in the late the nineteeenth century. Moreover, our findings are also consistent with those of several studies that rely on alternative methodologies to establish capital mobility.[20] We conclude that, when heterogeneity is incorporated into a panel framework robust to $I(1)$ innovations in finite samples, the FH puzzle virtually disappears.

There remains the question of how can one reconcile the contrasting inference from the traditional CS approach and the MG procedure. We conjecture

[17] The CS estimator is particularly sensitive to outliers as the well known case of Luxembourg illustrates.

[18] For a heterogeneous DGP the conditional variance of the CS residuals comprises a term that is proportional to $\overline{x}_i^2$. For large T this may become important and appropriate standard errors (0.093) can be obtained using White's correction.

[19] Defining an outlier as an MG coefficient more than two standard deviations from the mean $Z(\hat{\beta}_i) = |\hat{\beta}_i - \overline{\beta}|/\sigma(\hat{\beta}_i)$, none was found.

[20] For instance, Ho (2002) finds evidence of a low saving-investment association using Kao and Chiang's (2001) DOLS model. Similarly, De Vita and Abbott (2002) find a lowering of the saving-investment association to 0.20 for the US in recent decades using an ARDL model.

that country heterogeneity plays a significant role in explaining the relatively higher CS estimates.[21] While the MG estimator allows for country-specific intercepts and slopes, the former imposes homogeneity. Using the properties of OLS, it can be shown that $\hat{\beta}^{CS} = \hat{\beta}^{MG} + \delta$ where δ is the coefficient of a regression of $x_i'(\hat{\beta}_i - \hat{\beta}^{MG})$ on x_i'.[22] Our results thus bear out Corbin's (2001) and Taylor's (2002) argument that country-specific effects can be consequential in analysing the saving-investment association. Moreover, there remain questions surrounding the behaviour of the CS estimator for nonstationary panels of small sample dimensions.

5 Conclusions

Despite ongoing financial liberalization in recent decades, Feldstein-Horioka regressions have continued to indicate a high saving-investment association implying low long run capital mobility. This has been explained by the fact that a country's intertemporal budget constraint implies current account stationarity or that its saving and investment rates should cointegrate. However such behaviour may not be observed in finite samples since the current account could be subject to persistent shocks. Thus the puzzle is reconsidered in a nonstationary panel framework able to deal with this. The methodology employed builds on the asymptotic findings in Phillips and Moon (1999) and Kao (1999) and the related finite sample evidence in Coakley et al. (2001). These demonstrate that it is possible to overcome the time series nonsense regression problem by using panels. The intuition is that the noise that swamps the pure time series signal representing the long run relationship is attenuated by adding cross-section information.

The significant cross section slope estimate for 12 OECD economies 1980I-2000IV still conforms with the "two thirds" rule implying low capital mobility. However the mean group slope estimate is only 0.33 and, more importantly, insignificantly different from zero. This supports the hypothesis of long run capital mobility. By decisively overturning the verdict for recent

[21] It may also be the case that these estimators provide different measures of long run capital mobility. Whereas the mean group estimator gives an average long run measure, the cross section estimator may be yielding a measure of long run average capital mobility. The nature and implications of this issue warrant further research.

[22] This can be obtained by summing over T the estimated time series relation $y_{it} = x_{it}'\hat{\beta}_i + \hat{\varepsilon}_{it}$, where x_{it} and $\hat{\beta}_i$ are $k \times 1$ vectors, and noting that $\hat{\varepsilon}_i = T^{-1} \sum_t \hat{\varepsilon}_{it} = 0$.

decades in most of the extant literature, it leads us to conclude that the FH puzzle is not as bad as you think. In future work it would be interesting to test whether this result is robust when one extends the panel to allow for longer time spans and larger group dimensions.

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Table 1
Unit root and cointegration tests 1980I-2000IV

Country	Augmented Dickey-Fuller statistic		Augmented EG statistic
	s_t	i_t	
Australia	-3.0097(4T)	-3.1887*(2)	-2.9986(2)
Canada	-0.7642(8)	-0.9039(7)	-1.6980(5)
Finland	0.4060(8T)	-0.6237(5T)	-2.0289(5)
France	-2.7481(3T)	-2.3396(3)	-2.0731(7)
Italy	-2.0566(7)	-3.0223*(1)	-2.8909(1)
Japan	-1.7230(5)	-2.6712(7)	-3.2615(7T)
Netherlands	-4.2931*(3T)	-2.4534(7T)	-2.3727(7)
Norway	-1.2647(7T)	-1.2647(4)	-1.8350(7)
Spain	-1.8203(7)	-1.4042(8)	-1.2870(8)
Switzerland	-2.9652(8)	-3.0797(5T)	-3.3273(5T)
UK	-2.3327(8)	-2.0370(3)	-1.8518(3)
US	-3.6599(5T)	-1.5999(4T)	-2.9151(5)

*Denotes significant at the 5% level using Davidson and MacKinnon (1993) asymptotic critical values for the two tests.

Notes: Numbers in parenthesis refer to augmentation lag and T refers to a time trend.

Table 2
FH slope estimates 1980I-2000IV

	CS approach	MG approach
$\hat{\beta}$	0.6762	0.3276
$se(\hat{\beta})$	(0.1095)	(0.1765)
t-ratio (β=1)	-2.9572	-3.8091
t-ratio (β=0)	6.1744	1.8557*

*Non-rejection at the 5% significance level.

Table 3

Dispersion of country estimates $\hat{\beta}_i$

Country	$\hat{\beta}_i$	$Z(\hat{\beta}_i) = \frac{\lvert\hat{\beta}_i - \beta\rvert}{\sigma(\hat{\beta}_i)}$
Australia	0.5309	0.3326
Canada	0.7333	0.6635
Finland	-0.2295	0.9109
France	0.5794	0.4118
Italy	0.0874	0.3928
Japan	1.2120	1.4463
Netherlands	0.1134	0.3503
Norway	-0.5590	1.4499
Spain	-0.6924	1.6680
Switzerland	0.7965	0.7669
UK	0.2588	0.1124
US	1.1000	1.2632

Notes: The standard deviation is $\sigma(\hat{\beta}_i) = 0.6115$