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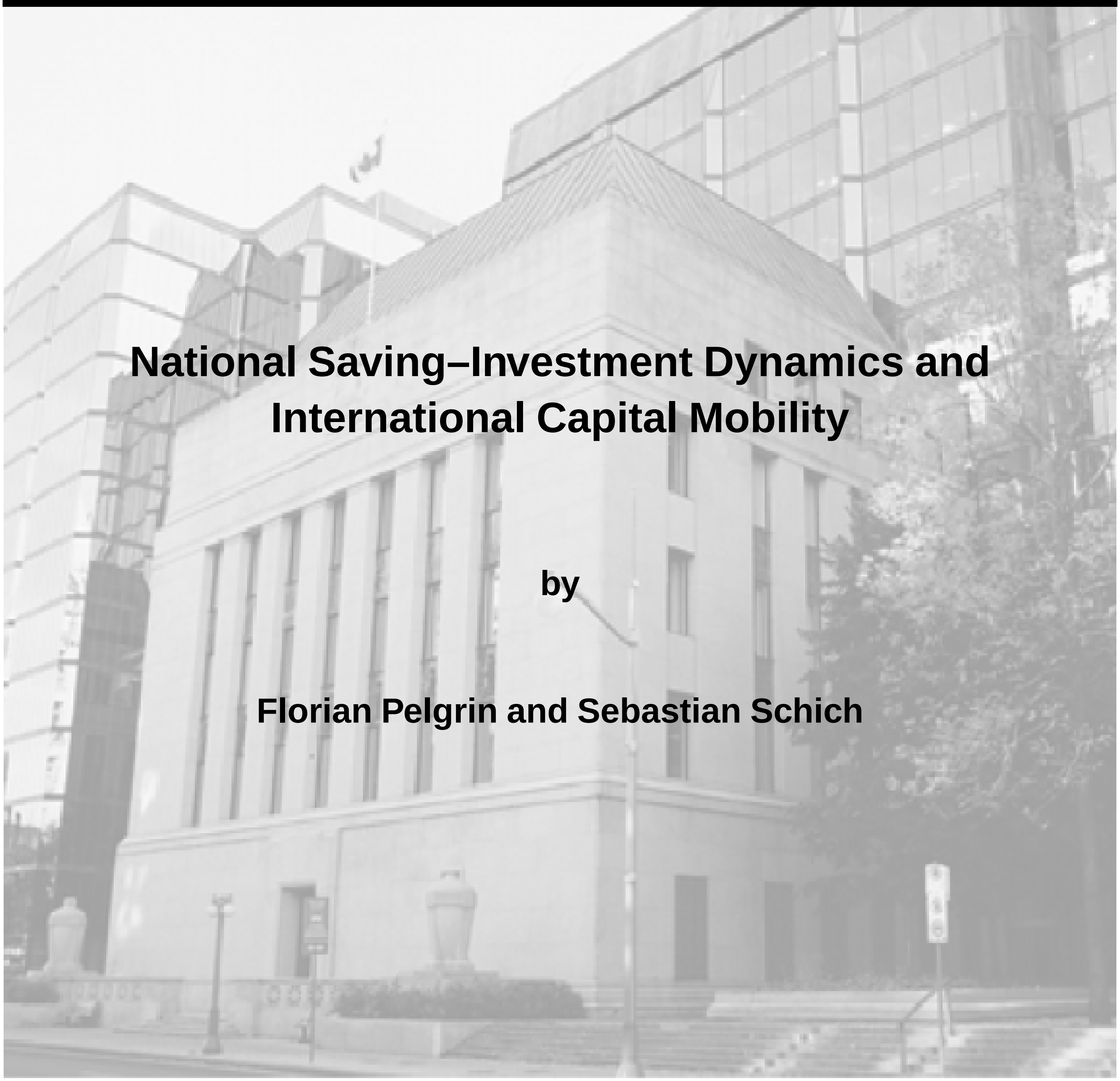

National Saving–Investment Dynamics and International Capital Mobility

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Abstract

The authors analyze the dynamics of national saving–investment relationships to determine the degree of international capital mobility. Following Coakley and Kulasi (1997), the authors interpret the close relationship between national saving and investment in the long run as reflecting a solvency constraint, rather than as evidence of limited capital mobility (Feldstein and Horioka 1980). As in Jansen (1996, 1998), the authors also examine the short-term saving–investment relationship, especially the speed at which the variables return to the long-run equilibrium relationship once they have deviated from it. The ease with which a country can borrow or lend and run current account imbalances in the short run, before it has to ultimately reverse the transaction at some future date to satisfy the solvency constraint, is interpreted as being positively related to the degree of international capital mobility. Extending the approach by Jansen, the authors apply panel error-correction techniques to data for 20 OECD countries from 1960 to 1999, and find that saving and investment display a long-run relationship that is consistent with the interpretation that a long-run solvency constraint is binding for each country. Furthermore, capital mobility has increased over time.

JEL classification: C23, F31
Bank classification: International topics

Résumé

Les auteurs étudient la dynamique de la relation entre l’épargne et l’investissement national pour déterminer le degré de mobilité internationale du capital. Suivant les traces de Coakley et Kulasi (1997), ils interprètent cette étroite relation de long terme comme reflétant une contrainte de solvabilité plutôt que le degré de mobilité du capital (Feldstein et Horioka, 1980). À l’instar de Jansen (1996 et 1998), ils examinent également la relation de court terme entre l’épargne et l’investissement et, plus particulièrement, la vitesse à laquelle les variables reviennent à leur valeur d’équilibre. La facilité avec laquelle un pays peut emprunter ou prêter à court terme et enregistrer un déséquilibre persistant de sa balance courante, avant de devoir inverser le flux de capitaux pour satisfaire à la contrainte de solvabilité, est interprétée comme une fonction directe de la mobilité du capital. Dans la foulée des travaux de Jansen, les auteurs recourent aux techniques de l’économétrie des panels pour estimer sur la période 1960-1999 des modèles à correction d’erreurs portant sur un échantillon de 20 pays de l’OCDE. La relation de long terme qu’ils observent entre l’épargne et l’investissement est conforme à leur hypothèse voulant que chaque pays ait à respecter une contrainte de solvabilité en longue période. Ils montrent en outre que la mobilité du capital a augmenté au cours de la période considérée.

Classification JEL : C23, F31
Classification de la Banque : Questions internationales

1. Introduction

Feldstein and Horioka (FH) (1980) identify a close cross-section association between period-average data on annual national saving and investment rates for a sample of 16 OECD economies from 1960 to 1974 and interpret it as evidence of low international capital mobility.[1] The authors argue that a *systematic* relationship between national saving and investment would not be expected if each country faced a large international capital market to which it supplied its national saving, or from which it obtained its means for investment purposes. Obtaining a saving rate coefficient that differs significantly from the benchmark value of zero in an investment rate regression implies a systematic relationship between saving and investment, and would thus be inconsistent with the view that there is sufficiently high capital mobility. A number of studies have re-estimated this relationship. The estimated saving–investment association often becomes weaker as more recent data are included, but it nevertheless remains significantly different from zero.

FH's results have generated a large literature on the saving–investment relationship.[2] Many alternative theoretical explanations have been proposed for the observed high and significant coefficient in a regression of investment on saving rates. For example, Obstfeld (1986) suggests that this reflects low international capital mobility as a result of information constraints and the lack of international enforceability of contracts. Bayoumi (1990) points out that a saving–investment correlation may reflect the fact that the government uses fiscal and monetary policies to target the current account. Harberger (1980) and Murphy (1984) argue that FH's result reflects a large country bias rather than low capital mobility. Obstfeld (1986, 1995) also suggests that the estimated saving–investment relationship may be indicative of a common effect or a third factor, such as productivity shocks. Simple general-equilibrium models have been constructed that reconcile high international capital mobility with a close correlation between national saving and investment rates by introducing productivity shocks in their models (Cardia 1991; Baxter and Crucini 1993).

Although some recent studies cast doubt on whether the empirical saving–investment relationship is indeed that close (Ho 2002), many others confirm the existence of a close relationship using non-stationary time-series analysis. They have interpreted this finding, however, as indicating

[1] See also Feldstein (1983).

[2] For a survey, see Coakley, Kulasi, and Smith (1998), and Taylor (1996).

a solvency constraint, rather than low capital mobility. Specifically, since the intertemporal budget constraint of an open economy should not allow countries to run current account deficits indefinitely (Sinn 1992), there must be a long-run relationship that ties national saving and investment together. Consequently, identifying empirically that saving and investment cointegrate is uninformative about capital mobility, because it reflects only the solvency constraint (Coakley, Kulasi, and Smith 1996, henceforth CKS; Coakley and Kulasi 1997, henceforth CK). Nevertheless, Jansen (1996, 1998) suggests that the short- and long-run dynamics of the saving–investment relationship could be used jointly to detect the degree of international capital mobility. Provided that a long-run cointegration relationship between saving and investment exists – reflecting that a solvency constraint is binding in the long run – identifying that the cointegrating vector is different from (1, -1)' could then be interpreted as evidence of capital mobility. The current account would be a non-stationary variable. This variable, however, could be non-stationary only if capital mobility were high.[3]

We follow CKS, CK, and Jansen, and consider that the short-run saving–investment dynamics may reflect fundamentally different phenomena from those that determine the long-run saving–investment relationship. We interpret the close long-run saving–investment relationship as reflecting a solvency constraint and focus on the short-term saving–investment relationship to assess the degree of capital mobility, since the latter is related to the process of an economy`s short-run adjustment to shocks.[4] In particular, we interpret the "speed" of an economy`s adjustment to shocks as another measure of the extent of capital mobility, based on an argument similar to that by Moreno (1997) and Jansen (1996, 1998). Moreno uses a VAR approach and suggests that inferences about capital mobility can be made from the extent of the divergence of short-run (dynamic) responses of saving and investment to shocks, particularly the "speed" at which the variables return to this long-run equilibrium relationship once they have deviated from it. Jansen (1996, 1998) interprets the speed of adjustment as being related to the extent of capital mobility, provided that a long-term saving–

[3] Jansen considers three ways to detect capital mobility. One way focuses on the long-run saving–investment relationship: he suggests that non-rejection of the hypothesis that the long-run saving–investment correlation is zero constitutes evidence in favour of international capital mobility. Our estimates, however, suggest that saving and investment are correlated in the long run. This, in turn, suggests that it is useful to focus on the interpretation of the long-run saving investment relationship proposed by CK and CKS; namely, that this relationship is mainly informative about the extent to which a solvency constraint is binding. In this context, Jansen also finds that the results obtained from the measure of capital mobility based on short-term saving–investment correlations differ from those based on the long-term saving–investment correlations.

[4] Clearly, other interpretations may also be valid, because the saving–investment regression is neither a structural relationship nor a reduced-form solution.

investment cointegration relationship exists that is significantly different from (1, -1). Using similar arguments, we interpret the estimates of the error-correction coefficient, l, from an error-correction model as an indicator of capital mobility. To see why, consider that, in the long run, saving and investment cointegrate because of the solvency constraint. If l were equal to one, adjustment to the long-run equilibrium would be immediate; in other words, the solvency constraint would be binding already in the short run. Capital mobility, however, allows saving and investment to deviate temporarily from each other. And the higher the capital mobility, the more prolonged such episodes could be, during which investment rates deviate from their long-run equilibrium levels, which are determined by the solvency constraint. Capital mobility is thus defined as the ease with which a country can borrow or lend to run prolonged current account imbalances in the short to medium term before it has to ultimately reverse the transaction at some future date. This is measured by the error-correction coefficient, l, with a lower l being consistent with higher capital mobility. Clearly, if l is very low and insignificant, this would suggest that the use of an error-correction approach is not appropriate. But, as long as l is significantly different from zero, the use of an error-correction approach is appropriate and we could interpret the estimated l as a measure of the extent of capital mobility.

An important difference between our estimation approach and the one used in Jansen (1996, 1998) is that ours obtains estimates directly from a panel of time-series and cross-section data, rather than from time-series estimates for individual countries. To our knowledge, our paper uses for the first time in the literature a panel error-correction model (Pesaran, Shin, and Smith 1999) to infer the degree of international capital mobility from the saving–investment dynamics. By using a panel of time-series and cross-section data and testing, simultaneously, the implications of the theoretical considerations on more than one country, increased power and more precise estimates are obtained. Moreover, it appears preferable to measure capital mobility in a large number of countries jointly, and thus take into account relations between them, rather than focus on individual countries when computing estimates regarding saving–investment dynamics. Some previous studies (Corbin 2001; Coakley, Fuertes, and Spagnolo 2001) have explored the panel nature of the data in a static context, but our approach differs in that we take into account possible dynamics. This is clearly useful because saving and investment are dynamic processes and static specifications are unlikely to capture the essential features of such processes. As well, our approach takes into account possible heterogeneity between countries, whereas many other studies of the saving–

investment relationship simply assume it to be equal across countries when pooling them in their data set. Our approach is also useful because the saving–investment relationship and the speed of adjustment to shocks can be expected to differ across countries, reflecting differences in their sizes, structures, cyclical positions, and policies. Indeed, the results in Jansen (1996, 1998) suggest that the short-term saving–investment relationship varies more substantially across countries than it does over time. We explicitly address the issue of heterogeneity by using three different estimation techniques that differ in their assumptions regarding homogeneity/heterogeneity in both short-term and long-term relationships. Specifically, in addition to the *dynamic fixed-effects* (DFE) *estimator*, we use *pooled mean group estimates* (PMG) and *mean group estimates* (MGE) . These three panel error-correction approaches differ with respect to their homogeneity assumptions. While the DFE imposes both the long-run and short-run coefficients to be the same between cross-section units of the panel, MGE allow both to differ across countries. PMG take a middle position in that they allow the short-run adjustment to differ across countries, while assuming that the long-run relationship is similar across countries. The data consist of saving and investment, as a percentage of GDP, for 20 OECD countries from 1960 to 1999 (from the *OECD National Accounts Database*).

Our empirical results are as follows. The estimated coefficient for the level of the saving rate is significantly different from zero, which suggests that a long-run cointegration relationship exists. This is consistent with the interpretation that a solvency constraint is binding in the long run. For the short-term adjustment, we find that the parameter estimated for the error-correction term is always highly significant, thus providing support for our choice of an error-correction formulation. The parameter estimated for the error-correction term (i.e., the speed of adjustment to the long-run equilibrium) varies with the sample period considered: using either a constant moving window of 30 years or a window of increasing size, the estimated l becomes smaller in absolute terms as more recent data are included in the sample. This is consistent with the hypothesis that capital has become more mobile over time. These basic results are robust with respect to the assumptions made concerning the homogeneity of the estimated parameters; i.e., whether we use DFE, MGE, or PMG. Nevertheless, our results suggest that it is important to allow for heterogeneity in the short-term relationship, since adjustments to shocks do differ substantially across countries. We suggest that the PMG and MGE are flexible enough to take that heterogeneity into account, while gaining efficiency by imposing some constraints regarding the long-term relationship.

The remainder of this paper is organized as follows. In section 2, the econometric methodology is explained. The results are presented in section 3. Section 4 concludes.

2. Econometric Methodology

We use a panel error-correction approach to analyze the long-run relationship separately from the short-run adjustment, estimating long- and short-run effects jointly from a general autoregressive distributed-lag (ARDL) model suggested by Pesaran, Shin, and Smith (1996).[5] This approach appears logical, since we take intertemporal general-equilibrium models as our theoretical frame (Jansen 1996). Our error-correction approach synthesizes various models; it offers different possibilities to assess the degree of international capital mobility, thus encompassing the different specifications previously used in the literature. The model is an extension of the time-series analysis and pooling approach used by Jansen (1996, 1998). It drops the homogeneity assumption implied by the pooling approach in the latter, combining both cross-section and time-series analysis to more fully exploit the cross-country variation in the data.[6] A benefit of our error-correction approach is that such a technique could capture the possibility that the short-run and long-run saving–investment correlations reflect phenomena that are fundamentally different from each other. For example, the long-run saving rate coefficient is expected to be driven mainly by the intertemporal budget constraint, whereas the short-run dynamics are likely to reflect short-run adjustments of the economy to shocks. The latter are determined by factors such as the structure, size, policy mix, stage of development, and cyclical position of a country. Depending on these factors, in the short-term, some countries will be net importers, while others will be net exporters of capital. In the long run, however, the solvency constraint implies a common constraint for all countries. This is taken into account by imposing similar constraints just on the long-run saving–investment relationship across countries.

[5] The main requirement for this approach is that a long-run relationship exist among the variables of interest (Pesaran, Shin, and Smith 1999). This can be verified by identifying a significant and negative coefficient estimate for the error-correction term. In addition, the regressors must be strictly exogenous and the residuals serially uncorrelated. While, strictly speaking, a Sargan test would be required to verify these assumptions, here we follow Pesaran, Shin, and Smith (1999) and test for robustness of the results by augmenting the lag length. Furthermore, the approach assumes that the saving rate is independent of past values of the investment rate, thereby ensuring that the long-run relationship is unique. This is a standard assumption also used in panel cointegration analysis.

Compared with traditional cross-section, time-averaging, or fixed-effects models, our approach has the following advantages. First, it allows us to distinguish between the short-term and long-term relationships. Second, the information available in the data can be more fully exploited, since the method does not rely on averaged data. Averaging data induces a loss of information. Third, the *dynamic* relationship between investment and saving rates can be analyzed. Fourth, this approach allows us to account for several types of heterogeneity in the saving–investment relationship across countries. To account for heterogeneity, Corbin (2001) uses the "within estimator" to correct for the possible bias due to the existence of specific countries effects assumed to be correlated with the investment rate. Corbin also uses the random-effects model to introduce a specific unobservable country effect in the error term. By contrast, our approach allows us to account for a larger variety of types of heterogeneity by including fixed effects and allowing the short-run as well as the error variances to differ across countries.

We use three different techniques: the DFE, MGE and PMG estimators. The methods differ in the extent to which they allow for cross-country heterogeneity of parameter estimates. At one extreme, the fully heterogeneous coefficient model, MGE, imposes no cross-country constraints and is estimated on a country-by-country basis. Coefficient estimates are obtained as the unweighted mean of the estimated coefficient for the individual countries. At the other extreme is the fully homogeneous coefficient model, the DFE estimator. It imposes the equality of both the saving rate coefficient and error variances, allowing only the intercepts to differ across countries. The PMG estimator can be interpreted as an intermediate procedure between the DFE and MGE, since it involves a mixture of pooling and averaging: it imposes the equality of the long-run saving rate coefficient, but allows the short-run coefficients to vary across countries.[7] The choice among these estimators is a trade-off between consistency and efficiency. In effect, estimators that impose restrictions dominate the heterogeneous models in terms of efficiency if the restrictions are valid. In particular, if the long-run coefficients are equal across countries, then the PMG will be consistent and efficient, whereas the MGE will only be consistent. If the long-run restrictions are wrongly imposed, however, the PMG estimates will be inconsistent. In that case, imposing invalid parameter

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[6] Jansen (1996, 1998) also considers one combined cross-section and time-series estimator, as he calculates the averages of the individual country results. This estimator is similar to the mean group estimator used here.

[7] Coefficients are estimated by the maximum-likelihood method after pooling, because of the homogeneity restrictions on the long-run coefficient. Based on individual results, the average across groups yields the mean of the estimated error-correction coefficient.

homogeneity in dynamic models would typically lead to downward-biased estimates of the speed of adjustment (Robertson and Symons 1992; Pesaran and Smith 1995). At the same time, the MGE are more sensitive to extreme individual results and are consistent only for large N and T. For small T (and any N), the familiar lagged dependent variable bias may cause the estimates of the speed of adjustment to underestimate their true values.[8] By contrast, in the case of the PMG estimator, the downward lagged dependent variable bias is at least partly offset by an upward heterogeneity bias.

Overall, it is difficult to say a priori which one of the three approaches is more appropriate in the context of the present sample; nevertheless, three considerations seem important. First, economic theory provides some guidance on long-run parameters, but is typically silent on short-term dynamics and the specific nature of the adjustment process. Second, the OECD countries form a relatively homogeneous group of countries, especially in terms of the extent to which they have liberalized their capital accounts. This would suggest that the long-run relationship between saving and investment rates is similar across these countries; for example, as a result of mobile capital eliminating arbitrage opportunities. Nevertheless, some impediments remain to capital mobility and countries differ substantially in terms of size and economic structure, so that, even within the group of relatively homogeneous OECD countries, differences in the long-run relationship between saving and investment rates could persist. Third, unless the heterogeneity across countries in short-run responses and adjustments is taken into account, the estimation of the parameters of long-term relationships can be biased. Against the background of these considerations, we apply all three different approaches and evaluate their performance based on our assessment of the plausibility of results, as well as statistical tests of the homogeneity of error variances and of short- and long-run slope coefficients. Likelihood ratios and other standard tests, such as Hausman's test, can be easily carried out, since the PMG and DFE estimators are restricted versions of the set of individual equations considered by the mean group estimator.

[8] Pesaran and Zhao (1998) propose a bias-corrected MG estimator, which adjusts the long-run coefficients for this potential bias. Because we tend to give more weight to the PMG estimator anyway, the computation of the bias-corrected version of the mean group estimator was considered to be beyond the scope of this paper.

Considering a maximum lag of one, the three specifications written in error-correction form are as follows[9]:

$$\Delta invr_{i,t} = a_i + \hat{l}_{DFE}(invr_{i,t-1} - \hat{b}_{DFE} savr_{i,t}) + \hat{d}_{DFE} \Delta savr_{i,t} + e_{i,t} \text{ (DFE)}, \qquad (1)$$

$$\Delta invr_{i,t} = a_i + \hat{l}_{PMG}(invr_{i,t-1} - \hat{b}_{PMG} savr_{i,t}) + \hat{d}_{PMG} \Delta savr_{i,t} + e_{i,t} \text{ (PMG)}, \qquad (2)$$

and

$$\Delta invr_{i,t} = a_i + \hat{l}_{MGE}(invr_{i,t-1} - \hat{b}_{PMGE} savr_{i,t}) + \hat{d}_{MGE} \Delta savr_{i,t} + e_{i,t} \text{ (MGE)}, \qquad (3)$$

$$\forall i = 1,2,...,N, t = 1,2,...,T,$$

where $invr_{it}$ is the investment rate and $savr_{it}$ the saving rate in country i at time t: $\hat{b}_{DFE} = \hat{b}_i, \hat{l}_{DFE} = \hat{l}_i, \hat{d}_{DFE} = \hat{d}_i \ \forall i$, $\hat{b}_{PMG} = \hat{b}_i \ \forall i, \hat{l}_{PMG} = 1/N \sum_{i=1}^{N} \hat{l}_i, \hat{d}_{PMG} = 1/N \sum_{i=1}^{N} \hat{d}_i$, and $\hat{b}_{MGE} = 1/N \sum_{i=1}^{N} \hat{b}_i, \hat{l}_{MGE} = 1/N \sum_{i=1}^{N} \hat{l}_i$ .and $\hat{d}_{MGE} = 1/N \sum_{i=1}^{N} \hat{d}_i$.

Given equations (1) to (3), the steady-state equilibrium (in country i) can be defined in each case as follows:

$$a_i + \hat{l}_{DFE}(invr_i^* - \hat{b}_{DFE} savr_i^*) = 0 \qquad \text{(DFE)}, \qquad (4)$$

$$a_i + \hat{l}_{PMG}(invr_i^* - \hat{b}_{PMG} savr_i^*) = 0 \qquad \text{(PMG)}, \qquad (5)$$

and

$$a_i + \hat{l}_{MGE}(invr_i^* - \hat{b}_{MGE} savr_i^*) = 0 \qquad \text{(MGE)}, \qquad (6)$$

$$\forall i = 1,2,...,N, t = 1,2,...,T.$$

[9] The above error-correction equations are written in terms of current, rather than lagged, values of the exogenous variables, to allow an ARDL(1,0) model as a special case.

It is useful to compare our modelling approach with that of individual-country time-series adopted by Jansen (1996, 1998) and Ho (2002). In their approach, the steady-state equilibrium in country i is defined as:

$$\alpha_i + \hat{\lambda}_i(invr_i^* - \hat{\beta}_i savr_i^*) = 0 . \qquad (7)$$

Comparing equation (7) with equations (4) to (6), the latter combine cross-section estimates by imposing different constraints on the short-run, long-run, or speed of adjustment coefficients. This allows us to obtain more efficient estimates while, at the same time, exploiting more fully the variation in the data across countries. Neither our panel approach nor the individual time-series approach, however, takes into account possible interdependencies between the various countries in the sample. Although this is a common assumption even in panel data analysis, it may be somewhat restrictive.

Two situations regarding the estimates of $\hat{\beta}$ need to be distinguished. First, if $\beta = 1$, then the current account (expressed as a ratio of GDP) is equal to $\alpha_i / \hat{\lambda}_j$ in the long run according to equations (4) to (6), where j = DFE, PMG, and MGE. Similarly, if $\beta_i = 1$ in equation (7), the current account ratio equals α_i / λ_i in the long run. Thus, in either one of these two situations, the current account ratio would be a stationary variable, fluctuating around its long-run value, which is equal to either $\alpha_i / \hat{\lambda}_{DFE}$, $\alpha_i / \hat{\lambda}_{PMG}$, or $\alpha_i / \hat{\lambda}_{MGE}$, depending on the choice of estimator. Moreover, if the fixed effects are equal to zero (i.e., $\alpha_i = 0$ in either of these approaches), then the current account fluctuates around zero. Second, if $\beta \neq 1$, the current account in the long run is defined as

$$(savr - invr)_i^* = \alpha_i / \hat{\lambda}_{DFE} + (1 - \hat{\beta}_{DFE})\ savr_i^* \text{ (DFE)},$$

$$(savr - invr)_i^* = \alpha_i / \hat{\lambda}_{PMG} + (1 - \hat{\beta}_{PMG})\ savr_i^* \text{ (PMG)},$$

$$(savr - invr)_i^* = \alpha_i / \hat{\lambda}_{MGE} + (1 - \hat{\beta}_{MGE})\ savr_i^* \text{ (MGE), and}$$

$$(savr - invr)_i^* = \alpha_i \big/ \hat{\lambda}_i + (1 - \hat{\beta}_i) \ savr_i^* .$$

Thus, under these circumstances, the investment rate, *invr*, and the saving rate, *savr*, are not cointegrated with vector (1,-1)'. Instead, they are cointegrated with vector (1, - $\hat{\beta}_{DFE}$)', (1, - $\hat{\beta}_{PMG}$)', and (1, - $\hat{\beta}_{MGE}$)' in the cases of the DFE, PMG, and MGE estimator, respectively, and with vector (1, - β_i)' in the case of individual-country time-series analysis. The current account ratio is a non-stationary variable in all these cases. As the current account can be non-stationary only if there is capital mobility, the identification of a non-stationary current account ratio in country i can be seen as evidence in favour of capital mobility between that country and the other countries in the sample.[10]

Therefore, our approach encompasses various empirical approaches in the literature, providing more than one way to detect capital mobility. First, failure to reject $\lambda = 0$ means that the saving and investment rates are not cointegrated in the long run and this could be seen as evidence in favour of the hypothesis of high international capital mobility, according to the logic of the initial FH interpretation. Our empirical results, however, are inconsistent with this interpretation; they suggest a significant long-run saving–investment correlation and thus we will not pay more attention to this specific way of detecting capital mobility. Second, rejection of $\lambda = 0$ means that there is a long-run saving–investment relationship. Under these circumstances, two cases need to be distinguished. On the one hand, if the cointegrating vector is not statistically different from (1,-1)', no conclusion about capital mobility can be drawn. On the other hand, if the cointegrating vector is different from (1,-1)', the current account is non-stationary. This is evidence in favour of international capital mobility. Consequently, if there exists a long-run relationship and the cointegrating vector is different from (1, -1)', the error-correction coefficient is an indicator of capital mobility, which allows saving and investment to temporarily deviate from each other. And the higher capital mobility is, the more prolonged could be the episodes during which investment can deviate from its long-run equilibrium value, determined by the solvency constraint.

[10] Miller (1988), Gundlach and Sinn (1992), Argimon and Roldan (1994), and Shreffin and Woo (1990) examine the reasons for the non-stationarity of the current account. Jansen (1998) asserts that non-stationarity is consistent with a specific class of general-equilibrium models with infinitely lived agents.

Before explaining the results, it is useful to discuss the choice of normalization of variables. If non-stationary saving and investment levels are considered, then the current account is stationary and investment and saving are cointegrated. Alternatively, if the variables are normalized by GDP, then the current account to GDP is stationary. This is the standard approach chosen to avoid heteroscedasticity (see, e.g., Fieleke 1982 and Tobin 1983), and it is the approach we follow. Note, however, that the error term from the estimates will be an infinite sum of past shocks either to the saving rate (or to the level of saving in the alternative approach) or the investment rate (or to the level of investment). Thus, the standard approach runs the risk that the division by a common factor, in this paper GDP, may create an artificial correlation between the saving and investment rates, even when their levels are uncorrelated. For instance, an income shock could create such a correlation between saving and investment rates. Thus, we need to determine that the correlation between saving and investment rates is not the result of a common factor.

3. Empirical Application

3.1. Unit root and cointegration tests

Before reporting the results of our panel error-correction estimates, we focus on the long-term relationship using cointegration tests. Following the procedure suggested by CK and CKS, we test for cointegration of saving and investment rates to determine whether our data are consistent with the interpretation that a solvency constraint ensures that the two variables do not move apart over the long run. Having verified that saving and investment rates are variables that are each integrated of order one, we use both homogeneous and heterogeneous panel cointegration tests to check for cointegration between saving and investment rates. The homogeneous panel cointegration tests suggested by Pedroni (1995) and Kao (1999) assume that the saving rate coefficient is the same for all countries under the null hypothesis of no cointegration. As the left-hand column of Table 1 shows, almost all statistics reject the null hypothesis of no cointegration. Next, we apply two heterogeneous panel cointegration tests suggested by Pedroni (1999, 2000): the panel statistics test and the group-mean statistics tests. The first tests the null hypothesis of no cointegration, $\boldsymbol{H}_0 : r_i = 1\,\forall \boldsymbol{i}$, against the alternative hypothesis, $\boldsymbol{H}_a : r_i = r < 1\,\forall \boldsymbol{i}$; that is, it assumes a common value for r under the alternative hypothesis. Applying this test, the null hypothesis of no cointegration is rejected by each of the four panel statistics (see Table 1, which also contains more information on the different tests). The group-mean statistics tests have the same null hypothesis as the panel statistics tests, but their alternative hypothesis is $\boldsymbol{H_a} : \boldsymbol{r_i} < 1 \forall \boldsymbol{i}$. Formulated in terms of r_i rather than r, they allow heterogeneity across individuals and include homogeneity of coefficients across all country subsets as a special case. As Table 1 shows, all three group-mean statistics tests reject the null hypothesis of no cointegration. Thus, regardless of the specific assumption made regarding homogeneity of the long-run saving rate coefficient, our results suggest that saving and investment rates are cointegrated.[11] Thus, our results are similar to those of Oh et al. (1999), but differ from those of Ho (2002).[12] The cointegration of saving and investment rates provides one rationale for using an error-correction approach to analyze the dynamic relationship between these rates. In addition, the significance of the adjustment term in such a specification can be used to test the existence of a long-run relationship (Pesaran, Shin, and Smith 1996).

As explained earlier, it is important to test for the possible existence of a common effect in the saving–investment relationship (see also Obstfeld 1986). One common feature of the tests described above is the restriction that all cross-sections are independent.[13] However, this assumption of cross-sectional independence may be too restrictive. In particular, considering a world of just two countries, a current account deficit in one country must be financed by a current account surplus in another country. Therefore, the deficit in the first country is not strictly independent of the surplus in the second one. In addition, the assumption of independence rules out links between variables that reflect common shocks in the data. Such shocks would induce a strong cross-section correlation that would not disappear by aggregating the data. Recently, some authors have relaxed the assumption of independence (Bai and Ng 2003; Phillips and Sul 2002; and, for a survey, Moon and Perron 2003).

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[11] This conclusion is confirmed by the results of cointegration tests for each country individually, using the trace and the maximum eigenvalue statistic for the Johansen test. These tests suggest the existence of one long-run cointegrating relationship for almost all countries. The latter is remarkable, since Bodman (1995) and Oh et al. (1999), each using saving and investment data for the G-7, could not reject the null hypothesis of no cointegration between national saving and investment for any of the countries examined individually. We suspect that this reflects the lack of power of individual cointegration, as opposed to panel cointegration tests.

[12] Ho (2002) considers more than one estimator. While he finds evidence in favour of a cointegrating relationship between saving and investment rates according to one estimator (FMOLS), an alternative approach suggests, at best, some weak cointegration (DOLS). Based on the results of Kao and Chiang (2001), Ho concludes that the saving retention coefficient tends to be low and that saving and investment rates "are virtually uncorrelated."

[13] If this assumption is violated, the null hypothesis may be rejected too often (Banerjee, Marcellino, and Osbat 2001).

Most of them use an approximate linear factor structure to model the cross-sectional dependence and assume that, after having accounted for the common linear factors, the cross-sections are independent. To determine whether potential cross-country dependency may affect our results, we apply to our data the method suggested by Phillips and Sul (2002). First, the observations are orthogonalized and then the Maddala-Wu panel unit root test is applied to the error term. The *p*-values obtained for the saving and the investment rate are 0.972 and 0.999, respectively. Thus, as neither of the *p*-values is less than 0.05, the null hypothesis of a unit root is not rejected in either case. Consequently, saving and investment rates are I(1) irrespective of the presence of a common factor in each series. Nevertheless, we also test directly to determine whether a long-run cointegration relationship allows for the existence of a common factor, using the approach suggested by Bai and Ng (2003). The test confirms the evidence of a cointegration relationship between saving and investment rates. Thus, we conclude that the estimated saving–investment correlation is not primarily driven by a common factor.[14]

3.2. Main results of panel error-correction estimates

To determine the lag order of our ARDL model, we follow the standard practice (Pesaran, Shin, and Akiyama 1998; Pesaran, Shin, and Smith 1999; and Pesaran, Haque, and Sharma 2000) and develop a restricted version of the general ARDL(1,1) model that is data-acceptable and preserves degrees of freedom. Starting with an ARDL(1,1), we use the Schwarz-Bayesian criterion to endogenously determine the optimal lag length for each country separately. This criterion suggests that an ARDL(1,1) model should be used for ten countries, an ARDL(1,0) for eight countries, an ARDL (0,1) for one country, and a static model for another country.[15] On the basis of these results, we decide to keep the ARDL(1,1) specification for the panel of countries, but we also examine the sensitivity of our estimation results to the order of the ARDL model.[16] We find that the results are

[14] This approach would require a more careful analysis of the common factor. This is beyond the scope of our paper.

[15] An ARDL(1,1) specification is found for the United States, Japan, Germany, France, Austria, Belgium, Denmark, the Netherlands, New Zealand, and Norway; an ARDL(1,0) model for Italy, the United Kingdom, Canada, Finland, Ireland, Portugal, Spain, and Sweden; an ARDL(0,1) model for Australia; and a static model for Greece.

[16] Testing for sensitivity of results is useful, because homogeneity restrictions and dynamic specifications interact in a complex way. This implies that the optimal order for the country-specific estimates may not be the optimal order when cross-country homogeneity restrictions are imposed.

similar for variations in the lag order up to a maximum of two lags. We suggest that this reflects the fact that our sample size is sufficiently large with T equal to 40. In the following, we report the results using an ARDL(1,1) specification.

Table 2 shows the results of our panel error-correction estimates. Note that the error-correction coefficient is significant for each country at standard levels using the methodology of Banerjee et al. (1993), which confirms the existence of a long-term saving–investment cointegration relationship. This result is consistent with those obtained from the panel cointegration tests discussed in section 3.1. Table 2 shows that the long-run coefficient estimates of the saving rate are highly significantly different from zero. The estimates are close to one and range between 0.82 and 1.02, depending on the method and time period considered. They are always insignificantly different from one, according to the simple Wald test, except for the case of the DFE and PMG estimators for the sample from 1960 to 1990.

As an additional check to determine whether the estimates of the coefficient for the long-run relationship are equal to one, we conduct heterogeneous and homogeneous cointegration tests (with and without a common effect). These tests suggest that the long-run saving–investment correlation is indeed insignificantly different from one.[17] Thus, overall, our results are consistent with those of Coakley, Fuertes, and Spagnolo (2001). Using a non-stationary panel methodology (mean group procedure) for 12 OECD countries, they also identify a long-run saving–investment rate cointegration with a coefficient insignificantly different from one. Our results are also consistent with those of Taylor (2002), who shows that the current-account-to-GDP ratio is stationary for all countries in his sample, which suggests that every country obeys its long-run budget constraint. Our results, however, differ from those of Ho (2002), who finds a saving retention coefficient estimated by FMOLS (or DOLS) that is close to just 0.84 (0.47, respectively) in a sample of 20 OECD countries over the period 1961 to 1997. Our paper, however, differs from Ho (2002) both in terms of sample and estimation approach. To determine the extent to which the differences in results reflect the choice of the estimation method, we apply FMOLS and DOLS estimators to our data. The long-run saving rate coefficient is estimated to be equal to 1.03 and 0.70 over the period 1960–99,

[17] Results are not reported here, but are available on request from the contact author.

according to the FMOLS and DOLS estimators, respectively.[18] Thus, we conclude that the differences may in part reflect the choice of estimator.

Focusing on the short-term relationships, we find that the coefficient on the error-correction term is negative and always highly significantly different from zero, irrespective of the estimation technique used (Table 2). This confirms the interpretation that there exists a long-run relationship between the saving and investment rates and that the short run is driven by the extent of the gap between current and long-run equilibrium values. The estimated values of the error-correction term range between –0.30 and –0.35, depending on the technique used, which means that half of the adjustment is completed after about two years and three quarters and that the adjustment is completed after about four years. We obtain an error-correction estimate that is significantly different from zero and interpret this as consistent with capital mobility, because an estimate that is statistically insignificantly different from zero would imply immediate adjustment to any national saving or investment shocks, which would be required only in the case of the absence of capital mobility. Furthermore, we obtain low estimated values of the short-run coefficient of the changes in the saving rate. This, as explained earlier, signals some degree of capital mobility, since such a low coefficient estimate can be obtained only if capital is sufficiently mobile. In section 3.3, we analyze whether this indicator of capital mobility varies with the sample period considered.

While the highly significant estimates for both the long-term saving–investment relationship and the error-correction term are valid regardless of the estimation technique used, some differences in the specific results do depend on the technique used. As for which of the estimators is most appropriate, we single out three considerations. First, imposing no restrictions whatsoever (MGE), we find evidence of substantial differences in estimates across countries (Table 3). Against this background, imposing homogeneity regarding *both* short-run and long-run estimates (DFE) may be too restrictive. Second, for the other two approaches, the question essentially is whether the assumed homogeneity of the long-run relationships is justified. In this respect, the results of our cointegration tests provide us with little guidance. While we find clear evidence to support cointegration, our results are inconclusive regarding whether the cointegrating relationship is homogeneous or heterogeneous. Testing directly for homogeneity of long-run coefficients using a

[18] We also conduct estimations over the two periods, 1960-90 and 1970-99, and find that results are broadly similar. They are not reported here but are available on request.

likelihood-ratio test, we reject the hypothesis of equality at conventional levels. This test has been criticized, however, and Hausman's test suggested as an alternative (Pesaran, Shin, and Akiyama 1998). Using Hausman's test, we cannot reject the long-run restriction on the saving rate coefficient. Thus, taking the evidence of the various tests together, there is no strong case for either homogeneity or heterogeneity. Third, there are some differences in the long-run coefficient estimates, with some countries being, on average, capital importers and others capital exporters. For example, the United States is a capital importer with a saving rate coefficient below one, and Japan is a capital exporter with a coefficient that exceeds one. But while there may be positive or negative differences, on average, between saving and investment over the sample period, deficits and surpluses must be stationary. This constraint is identical to all countries, whether they are a net capital exporter or importer. We suggest that this requirement imposes homogeneity on the long-run saving–investment relationship across countries and that the efficiency gains offered by a panel data set are best exploited by PMG.

We also find some differences in the results for the short-run and long-run saving–investment coefficients and the estimated speed of adjustment. This confirms the results obtained by Jansen and appears intuitive. For example, the difference in the estimates of the speed of adjustment of the current account reflects typical differences in current account patterns. Some countries may run current account deficits for a long time, for country-specific structural reasons, before ultimately reversing the deficit and respecting the intertemporal budget constraint. Other countries may reverse current account deficits much more quickly or run surpluses. There is also heterogeneity in the estimates of the short-run saving–investment coefficients. This seems natural, since the response of the economy to shocks may differ across countries. The differences obtained in the estimated coefficients suggest that it is indeed useful not to constrain all coefficients to be the same a priori. This approach, however, is taken in traditional cross-section saving–investment regressions.

3.3. Further results and robustness analysis

We conduct various robustness tests in addition to the analysis described in the previous subsection. First, to determine whether our results reflect the inclusion of individual countries that substantially influence the overall estimates, we re-estimate our equations for all possible subsamples obtained by deleting one country at a time from the original sample. The estimated saving rate coefficients are shown in Figures 1 to 3, with estimates arranged in decreasing order and the names of countries used to identify that they are excluded from the sample. The figures show that the saving rate coefficients for the different subsamples do not vary much and are quite close to the estimate obtained on the basis of the full sample, regardless of the estimation technique used. Second, following Kim (2001), we test whether third factors such as country size or openness affect our estimation results. We divide our sample into three groups of countries, according to either their sample-average real GDP or the size of the non-traded sector, and perform separate panel regressions for each group.[19] The results (Table 4) show that, regarding country size, the estimated long-run relationship (i.e., the coefficients of the saving rate) are similar across different groups, and that they are always insignificantly different from one. This is consistent with the results reported in section 3.2. There is, however, some variation as far as the short-term relationships are concerned. In the large-country group, the estimated coefficient of the error- correction term is smaller (in absolute values) and the estimated coefficient of the change in the saving rate is larger than in the other two groups. While the difference in the coefficient estimates of the error-correction term is not significant, there is a significant difference in the results regarding the coefficient of the change in the saving rate: it is insignificant in the cases of medium and small countries, but significantly different from zero for large countries. This may explain the results of Baxter and Crucini (1993) and Tesar (1991), who find that accounting for country size effects causes the saving–investment relationship to become weaker or even insignificant. We suggest that there may indeed be a country size effect, but that it may affect only the short-term relationship. Consequently, if the short-term relationship is not estimated separately, this effect may influence the estimates of the long-term relationship. Our main results regarding the long-term relationship and the error-correction term are not affected by this size effect. In addition, as far as the openness of countries is concerned, the results are broadly similar,[20] except that the coefficient estimates for the change in the saving rate are significant not only for the group of large countries, but also for the group of small countries. The main results again do not differ across different groups. Thus, we conclude that neither single individual countries, nor the size of the GNP or of the non-traded sector, have significant effects on our results.

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[19] We also verify that the results are robust to small changes to the classification of countries in one of the groups.

[20] The long-run saving rate coefficients are statistically different from one for large countries in the case of the PMG estimator. According to our discussion in section 2, this could be interpreted either as evidence for

To determine whether the saving–investment relationship has undergone changes over time, we re-estimate the specification over various subsamples,[21] in two different ways. First, starting from a 30-year sample from 1960 to 1990, the sample period is extended one year at a time to include the more recent data up to 1999 (Figure 4). Second, the sample or window size is kept constant at 30 years, but the window is moved, so that the most recent one includes data from 1970 to 1999 (Figures 5 and 6). The results of the two approaches are broadly similar. If anything, the results from the first approach show less variation from year to year, possibly reflecting that the somewhat larger sample size allows more reliable estimates.[22] We do not suggest focusing on the year-to-year changes, but rather on the trend change in coefficient estimates. Indeed, both short- and long-run coefficients exhibit fairly smooth and mostly monotonic changes when the samples are changed. The figures show that the long-run coefficient estimates increase over time and become overall closer to one (shown on the right-hand scales), while the estimates of the error- correction term coefficient decrease in absolute terms (shown on the left-hand scales). The finding that the estimate of the long-run saving rate coefficient rises and becomes insignificantly different from one as the sample increases is interpreted as evidence that the solvency constraint becomes more relevant as the horizon increases. In other words, the solvency constraint is more likely to be binding over a period of 40 years than over 30 years. Regarding the short-run coefficients, we observe that they fall in absolute values. This is consistent with the hypothesis that capital mobility increased during the 1990s. Specifically, as more recent data are included, there is evidence that more prolonged deviations from long-run equilibrium relationship are feasible, as reflected in estimates of the error-correction term coefficients that are smaller in absolute values.

The fact that the long-run coefficient estimates increase over time as more data are included may be considered to conflict with the results of previous empirical studies, which tend to find lower coefficient estimates because they include more recent data. Our finding, however, is not inconsistent with the observation that the coefficient in FH regressions exhibits a downward trend. It

capital mobility or that the solvency constraint is less binding for large than for small countries. We favour the latter interpretation, but leave exploration of this issue for further work.

[21] One possible drawback of this approach is that results may be driven by sampling effects owing to increasing samples. The dimension of the panel, however, is large enough to avoid such an undesirable effect. Yet, it may affect individual results. We consider this effect to be negligible, because no outlier in the individual series has been identified.

[22] Wald tests for each year are not reported but are available on request.

may simply reflect the lack of dynamics in many previous studies. Indeed, the results shown in Table 5, obtained using standard estimators that do not account for dynamics, suggest that there is a decreasing coefficient for the relationship between the saving and investment rates. This finding has sometimes been interpreted as evidence for increasing capital mobility, but the framework used in this paper puts that interpretation in doubt.

4. Concluding Remarks

We have confirmed previous empirical results that identify a close relationship between saving and investment rates in the long run. This long-run relationship is considered to be consistent with a solvency constraint that is binding in the long run. Regarding the short-term adjustment, we have found that the parameter estimated for the error-correction term is always highly significant, which supports our choice of an error-correction formulation. Moreover, we have found that the parameter estimated for the error-correction term (i.e., the speed of adjustment to the long-run equilibrium) varies with the sample period considered. Using either a constant moving window of 30 years or a window of increasing size, the estimated ? becomes smaller in absolute terms as more recent data are included in the sample. This is consistent with the hypothesis that capital has become more mobile over time. These basic results are robust to our assumptions regarding the homogeneity of the estimated parameters (i.e., whether we use DFE, MGE, or PMG). Nevertheless, our results suggest that it is important to allow for heterogeneity in the short-term relationship, because adjustments to shocks do differ substantially across countries. We suggest that PMG and MGE are flexible enough to take that heterogeneity into account. Our results are indeed robust with respect to various alternative choices of the country sample. The exclusion of individual countries does not affect the results. Neither does the analysis of countries grouped according to either their size or openness. Moreover, to address the concerns expressed by Obstfeld (1986), we have applied modern tests to take into account the existence of a common effect and found that the estimated saving–investment relationship remains robust. We conclude that by jointly estimating the short-run and long-term relationship using a variety of techniques, and by applying those techniques to a panel of time-series data for many countries, some of the differences in results obtained by previous studies can be reconciled.

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Table 1. Panel Cointegration Tests

Homogeneous tests[1]		Heterogeneous tests[1]		
			Without common time dummy	With common time dummy
Pedroni (1995)[2]		*Pedroni (1999)*[5]		
PC1	-34.82***	*Panel statistics*		
PC2	-34.39***	Panel-v	8.17***	5.64***
		Panel-rho	-4.43***	-3.83***
		Panel-pp	-3.78***	-4.78***
		Panel-ADF	-4.01***	-5.07***
Kao (1999)[3]				
ADF	-6.40***	*Group-mean statistics*		
		Group-rho	-3.73***	-2.35***
		Group-pp	-4.13***	-4.50***
Kao (1999)[4]		Group-ADF	-4.82***	-5.23***
DF-rho	-13.69***			
DF-*t*	15.10			
DF-rho*	-22.60***			
DF-*t**	-6.22***			

Note: *** denotes rejection of null hypothesis of no cointegration at the 1 per cent significance level.

1. All cointegration tests are panel cointegration tests. Homogeneous tests assume equality of all coefficients in the null and alternative hypothesis, while heterogeneous tests allow them to differ across countries. All tests are left-hand side, except for the panel-v test (right-hand side). All statistics converge in distribution to standard normal distribution.
2. PC1 and PC2 are the non-parametric Phillips-Perron tests.
3. ADF is the Augmented Dickey-Fuller test.
4. DF is the Dickey-Fuller test. The DF-rho and DF-*t* statistics assume strict exogeneity of the regressors with respect to the errors and no autocorrelation. DF-rho* and DF-*t** statistics are based on endogenous regressors. The DF-rho (DF-rho*) and the DF-*t* (DF-*t**) are based on the standardized bias and the standard least-squares *t*-statistic for $\rho = 1$, respectively.
5. The panel statistics are based on estimators that pool along the within-dimension the autoregressive coefficient across countries for the unit root tests on the estimated residuals. The group-mean statistics are based on the average of the coefficients estimated for each country. The panel-v statitistic is a non-parametric variance ratio statistic. The panel-rho(group-rho) statistic is a non-parametric statistic similar to the Phillips-Perron rho-statistic. The panel-pp (group-pp) statistic is a non-parametric statistic similar to the Phillips-Perron *t*-statistic. The panel-adf (group-adf) statistic is a parametric statistic similar to the Dickey-Fuller *t*-statistic. The results shown are for an actual lag order of one. The results were qualitatively similar for higher lag orders.

Table 2. Panel Error-Correction Estimates of Saving–Investment Relationship

Estimators[1]	1960 – 99			1960 – 90			1970 – 99		
	DFE	PMG	MGE	DFE	PMG	MGE	DFE	PMG	MGE
Long-run coefficients									
Constant		0.01***	0.01		0.03***	0.02**		0.01***	0.01***
		(4.84)	(1.00)		(7.09)	(2.09)		(4.73)	(0.79)
Saving rate	0.93***	0.93***	1.01***	0.82***	0.84***	0.92***	0.94***	0.92***	1.03***
	(9.05)	(20.10)	(12.07)	(9.31)	(24.00)	(10.29)	(8.65)	(17.13)	(9.81)
Adjustment term									
Error-correction	-0.30***	-0.33***	-0.35***	-0.38***	-0.46***	-0.49***	-0.29***	-0.33***	-0.36***
term	(-8.33)	(-9.76)	(-9.90)	(-8.19)	(-8.64)	(-9.77)	(-8.45)	(-11.39)	(-11.31)
Short-run coefficients									
Δ(saving rate)	0.22**	0.25***	0.23***	0.17*	0.15	0.11	0.22**	0.22**	0.20**
	(2.22)	(2.91)	(2.65)	(1.77)	(1.59)	(1.15)	(2.06)	(2.53)	(2.29)
Memorandum item									
Hausman's test[2]		1.35			1.11			1.54	
[*p*-value]		[0.25]			[0.29]			[0.21]	

Notes: *t*-value in parentheses. *, **, *** denote the 10, 5, and 1 per cent significance level, respectively.

1. The dynamic fixed-effects estimator (DFE) assumes both long-run and short-run homogeneity of all coefficients. The pooled mean group estimator (PMG) allows short-run coefficients, including the adjustment term, and the variances of the error term to differ across countries, while the long-run saving rate coefficient is constrained to be the same. The mean group estimator (MGE) allows the long-run saving rate coefficient to differ between countries as well. For all estimators, an ARDL(1,1) specification is used, where the first number and second number in parentheses stand for the lag length of the lagged dependent and the explanatory variable, respectively. For ease of reference, we also report here the coefficient estimates for the saving rate according to the static fixed-effects estimator, which has been used in several previous papers: They are 0.66, 0.61, and 0.64 for the subperiods 1960 – 99, 1960 – 90, and 1970 – 99, respectively.

2. Hausman's test determines the validity of the assumption made for the long-run saving rate coefficient across OECD countries (i.e., comparing PMG and MGE estimation results).

Table 3. Individual Results of PMG and MGE Estimates of Saving–Investment Relationship, 1960 – 99

	Pooled mean group estimates (PMG)[1]			Mean group estimates (MGE)[1]			
	Adjustment term	Δ(saving rate)	Adjusted R-square	Saving rate	Adjustment term	Δ(saving rate)	Adjusted R-square
United States	-0.16*	0.65***	0.61	0.41**	-0.30***	0.62	0.65
	(-1.89)	(5.61)		(1.99)	(-2.67)	(0.12)	
Japan	-0.14	0.91***	0.67	1.46***	-0.18*	0.81***	0.69
	(-1.48)	(5.79)		(3.63)	(-1.78)	(4.73)	
Germany	-0.27***	0.45**	0.43	1.23***	-0.27**	0.42**	0.45
	(-2.59)	(2.30)		(3.88)	(-2.49)	(2.04)	
France	-0.33***	0.51**	0.47	0.95***	-0.32**	0.51**	0.47
	(-2.71)	(2.47)		(6.09)	(-2.42)	(2.35)	
Italy	-0.43***	0.12	0.30	0.93***	-0.43***	0.12	0.30
	(-3.42)	(0.43)		(4.92)	(-3.17)	(0.40)	
United Kingdom	-0.31***	0.12	0.21	0.79*	-0.32***	0.13	0.22
	(-2.92)	(0.72)		(1.76)	(-2.64)	(0.73)	
Canada	-0.45***	0.31*	0.60	0.84***	-0.49***	0.28*	0.60
	(-3.68)	(1.86)		(6.87)	(-3.49)	(1.64)	
Australia	-0.36***	0.44**	0.42	0.50***	-0.68***	0.29	0.53
	(-2.97)	(2.29)		(4.48)	(-4.13)	(1.53)	
Austria	-0.42***	0.52***	0.57	0.79***	-0.46***	0.52***	0.58
	(-3.07)	(2.70)		(6.22)	(-3.11)	(2.60)	
Belgium	-0.12*	0.55***	0.55	0.82*	-0.13	0.54***	0.55
	(-1.76)	(4.51)		(1.81)	(-1.16)	(3.64)	
Denmark	-0.26**	0.51**	0.32	0.78***	-0.27**	0.51**	0.33
	(-2.21)	(2.43)		(2.78)	(-2.16)	(2.31)	
Finland	-0.35***	0.22	0.43	1.77***	-0.31***	0.14	0.54
	(-2.78)	(0.87)		(3.91)	(-2.65)	(0.60)	
Greece	-0.60***	0.36**	0.77	0.87***	-0.60***	0.38**	0.77
	(-4.06)	(2.43)		(9.11)	(-3.86)	(2.38)	
Ireland	-0.25***	-0.14	0.16	1.60**	-0.21**	-0.18	0.19
	(-3.17)	(-0.62)		(1.99)	(-2.39)	(-0.73)	
Netherlands	-0.12*	0.46***	0.30	1.23	-0.11	0.45***	0.30
	(-1.67)	(3.00)		(1.51)	(-1.44)	(2.82)	
New Zealand	-0.64***	-0.65***	0.42	0.97***	-0.63***	-0.66***	0.43
	(-5.85)	(-3.14)		(4.87)	(-5.16)	(-2.98)	
Norway	-0.37***	-0.57***	0.27	1.38***	-0.34***	-0.61***	0.30
	(-4.26)	(-2.74)		(3.05)	(-3.58)	(-2.80)	
Portugal	-0.19**	0.19	0.22	0.43	-0.31**	0.18	0.26
	(-2.33)	(1.49)		(1.51)	(-2.43)	(1.32)	
Spain	-0.48***	-0.19	0.38	1.08***	-0.46***	-0.20	0.39
	(-4.64)	(-0.86)		(5.36)	(-4.20)	(-0.88)	
Sweden	-0.27**	0.33*	0.42	1.29***	-0.25**	0.32	0.45
	(-2.49)	(1.67)		(3.85)	(-2.22)	(1.57)	

Notes: The *t*-value is in parentheses. *, **, *** denote the 10, 5, and 1 per cent significance level, respectively. The value of the constant is not reported to save space.

1. The PMG constrains the coefficient of the saving rate to be the same for all countries. It is estimated to be equal to 0.93, with a *t*-value of 20.10. In the case of both PMG and MGE, an ARDL(1, 1) specification is used.

Table 4. Robustness of Estimates to Country Size and Openness, 1960-99[1]

Size of GNP[2]	Pooled mean group estimates (PMG)			Mean group estimates (MGE)		
	Large	Medium	Small	Large	Medium	Small
Long-run coefficients						
Constant	0.00	0.01	0.01	0.00	0.01	0.01
	(0.84)	(3.29)***	(3.16)***	(0.24)	(0.73)	(0.57)
Saving rate	0.98	0.90	0.95	0.96	1.00	1.06
	(9.66)***	(14.49)***	(9.79)***	(6.57)***	(9.62)***	(4.97)***
Adjustment term						
Error correction	-0.27	-0.35	-0.35	-0.30	-0.38	-0.37
	(-6.36)***	(-5.80)***	(-5.23)***	(-8.95)***	(-5.15)***	(-5.82)***
Short-run coefficients						
Δ(saving rate)	0.46	0.21	0.11	0.44	0.18	0.09
	(3.77)***	(1.55)	(0.58)	(3.94)***	(1.31)	(0.47)
Memorandum item						
Hausman's test[3]	0.04	1.57	0.34			
[*p*-value]	[0.85]	[0.21]	[0.56]			

Size of non-traded sector[4]	Pooled mean group estimates (PMG)			Mean group estimates (MGE)		
	Large	Medium	Small	Large	Medium	Small
Long-run coefficients						
Constant	0.03	0.01	0.01	0.03	-0.00	0.01
	(4.00)***	(2.07)**	(2.85)***	(1.41)	(-0.29)	(0.65)
Saving rate	0.71	0.98	0.83	0.80	1.12	1.04
	(7.79)***	(16.17)***	(8.04)***	(4.60)***	(10.47)***	(6.44)***
Adjustment term						
Error correction	-0.33	-0.39	-0.24	-0.39	-0.40	-0.24
	(-4.65)***	(-9.13)***	(-4.01)***	(-5.56)***	(-8.19)***	(-3.80)***
Short-run coefficients						
Δ(saving rate)	0.36	0.11	0.38	0.30	0.10	0.37
	(2.33)**	(0.79)	(2.97)***	(2.04)**	(0.71)	(2.68)***
Memorandum item						
Hausman's test[3]	0.42	2.54	2.99			
[*p*-value]	[0.52]	[0.11]	[0.08]			

Notes: The *t*-value is in parentheses. **, *** denote the 5 and 1 per cent significance level, respectively.

1. We perform separate panel regressions for each group of countries, similar to Kim (2001).

2. Groups are selected according to the size of the sample-average real GDP denominated in U.S. dollars. Accordingly, large countries are the United States, France, Japan, Italy, the UK, and Germany. Medium countries are Spain, Canada, Australia, Greece, the Netherlands, Belgium, Norway, and Sweden. Small countries are Portugal, Denmark, Austria, Finland, New Zealand, and Ireland.

3. Hausman's test determines the validity of the equality of the long-run saving rate coefficient across OECD countries by comparing the PMG and MGE estimates.

4. Groups are selected according to the size of the non-traded sector, which is approximated, following Wong (1990), by sample-average of the export minus the import over GNP ratio. Large countries are the United States, Japan, Australia, Portugal, Spain, and Italy. Medium countries include France, the UK, Canada, New Zealand, Germany, Finland, Greece, Norway, and Sweden. Small countries are Denmark, Austria, the Netherlands, Ireland, and Belgium.

Table 5: Estimation of the Saving Rate Coefficient with Static Panel Estimators

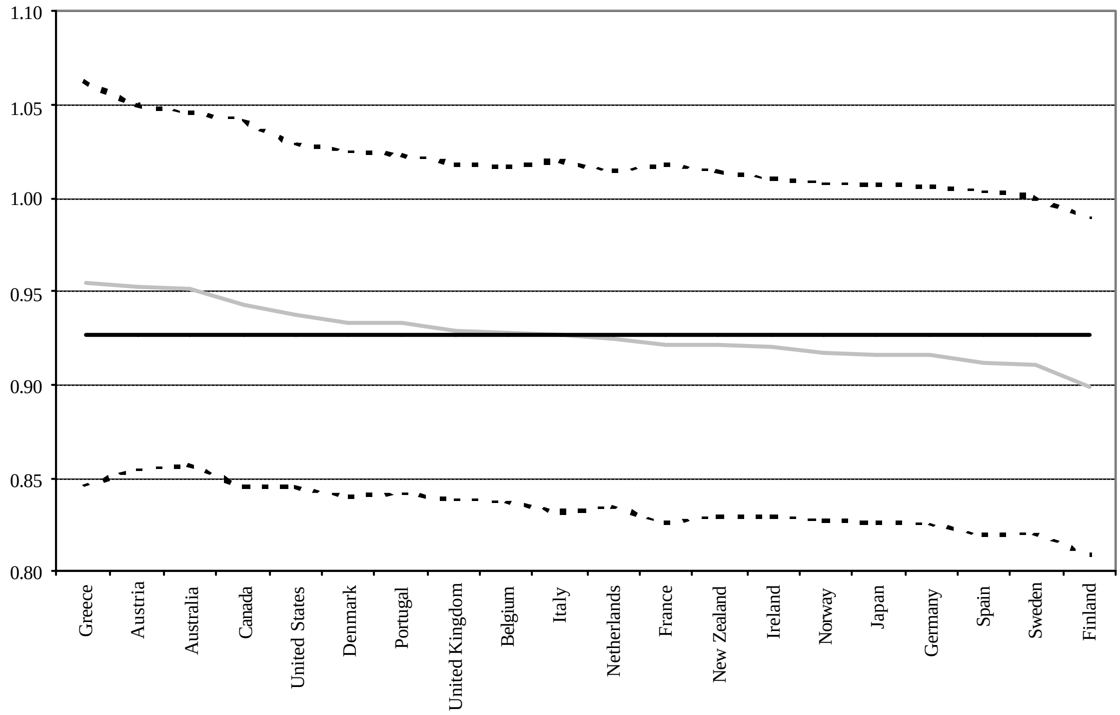
	1960 – 99	1960 – 69	1970 – 79	1980 – 89	1990 – 99
Between[1]	0.93***	0.99***	0.95***	0.85***	0.77***
	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)
Pooled[1]	0.95***	0.99***	0.94***	0.87***	0.77***
	(0.01)	(0.01)	(0.03)	(0.03)	(0.01)
Within[1]	0.93***	1.10***	0.83***	1.16***	0.80***
	(0.00)	(0.05)	(0.12)	(0.20)	(0.12)
Random effects[1]	0.97***	0.99***	0.95***	0.88***	0.77***
	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)
Memorandum item					
Hausman's test[2]	10.98	5.48	0.92	1.89	0.07
[*p*-value]	[0.00]	[0.02]	[0.34]	[0.17]	[0.80]

Note: White heteroscedasticity-consistent errors in parentheses. *** denote the 1 per cent significance level.

1. The between estimator is obtained from the average value of each country and therefore emphasizes the inter-country dimension. The pooled estimator assumes individual homogeneity as well as the temporal stability of the relation. The within estimator introduces heterogeneity through individual fixed effects and is calculated from the difference between the saving and investment ratio and the individual average of the variable. The random effects model introduces heterogeneity through a specific unobservable country effect in the error term.

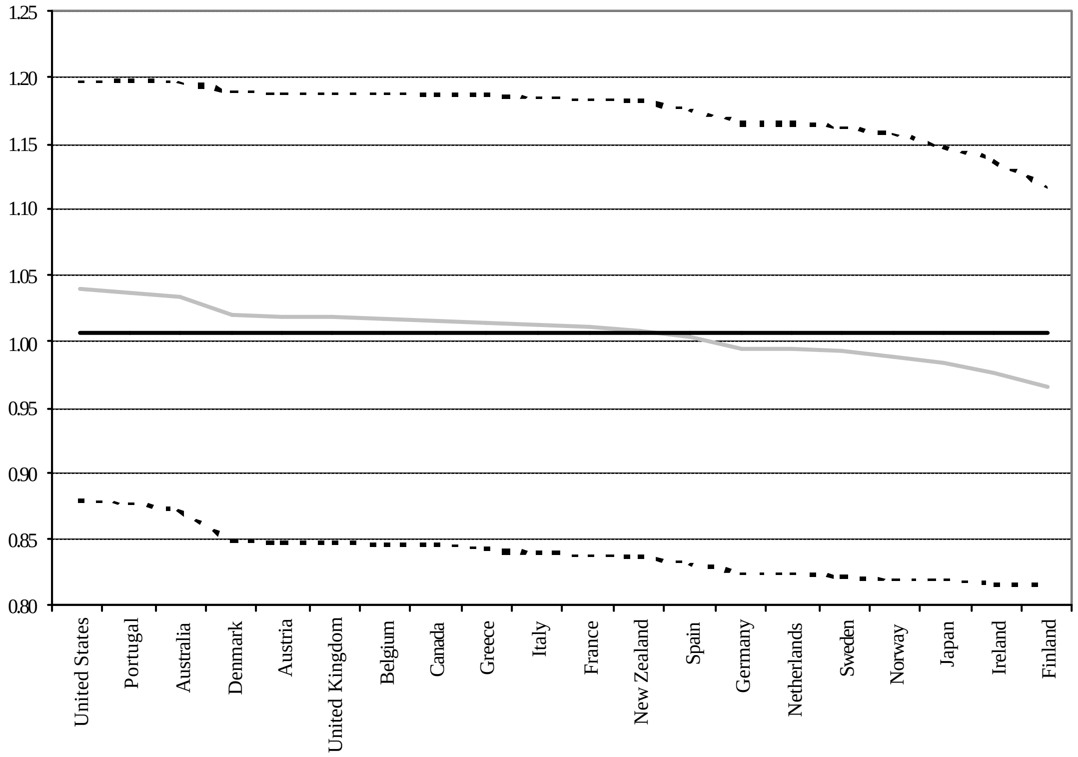
2. Hausman's test checks for fixed versus random effects.

Figure 1: Pooled Mean Group Estimates of Saving Rate for Different Samples



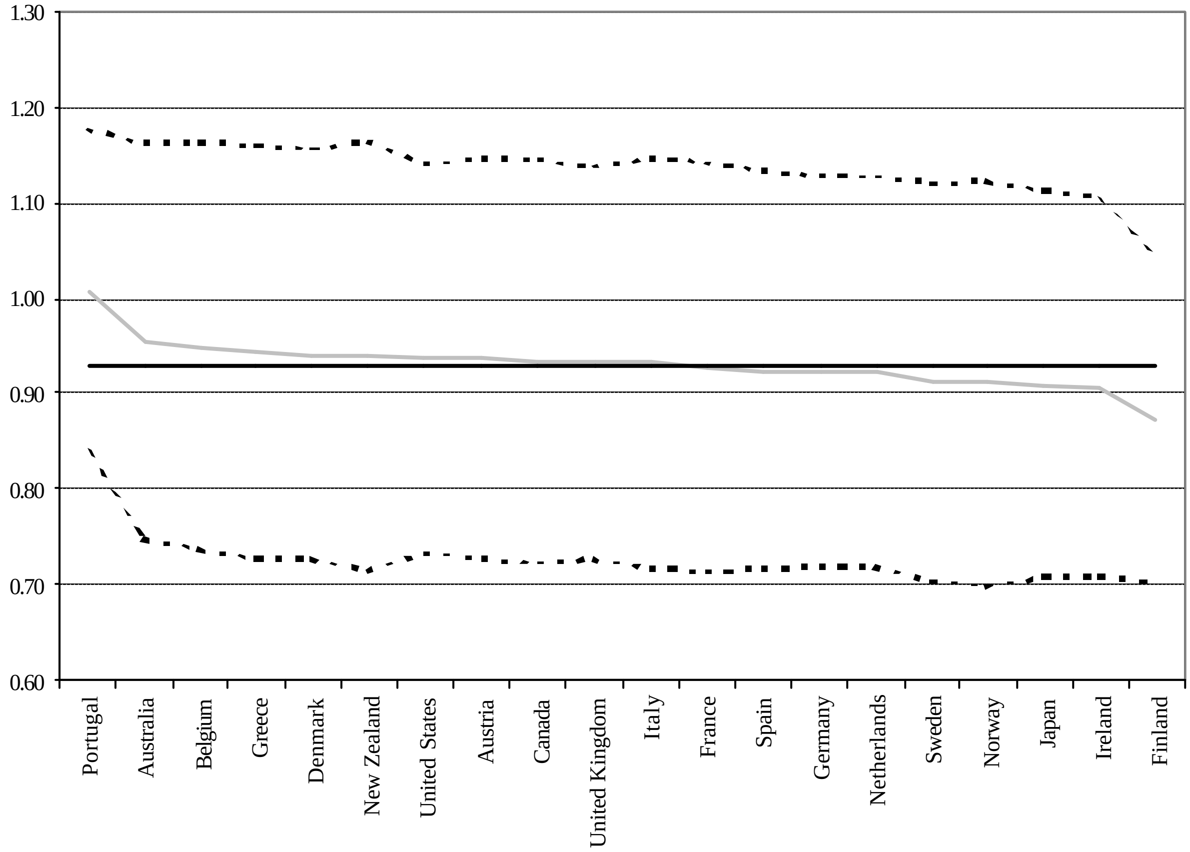
Note: Coefficient estimates and standard error bands according to PMG (95 per cent confidence interval around coefficient estimate) when one country is excluded at a time from the sample. The coefficient estimates are arranged in decreasing order.

Figure 2: Mean Group Estimates of Saving Rate for Different Samples



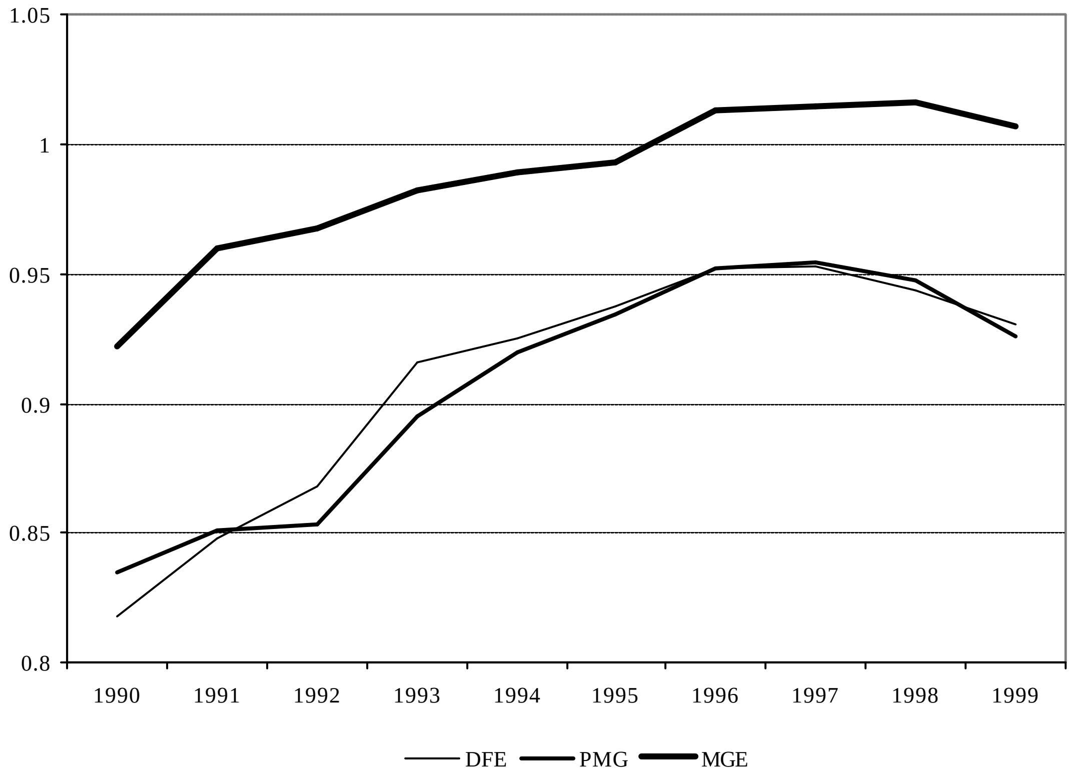
Note: Coefficient estimates and standard error bands according to MGE (95 per cent confidence interval around coefficient estimate) when one country is excluded at a time from the sample. The coefficient estimates are arranged in decreasing order.

Figure 3: Dynamic Fixed Effects Estimates of Saving Rate for Different Samples



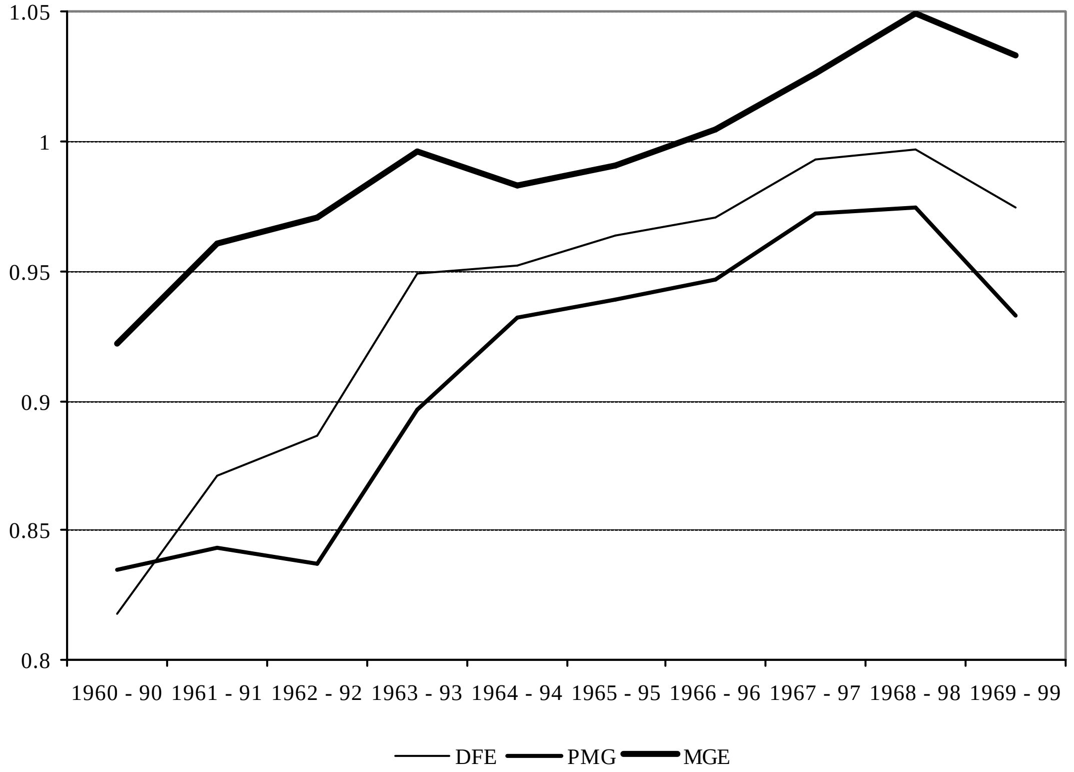
Note: Coefficient estimates and standard error bands according to DFE (95 per cent confidence interval around coefficient estimate) when one country is excluded at a time from the sample. The coefficient estimates are arranged in decreasing order.

Figure 4: Coefficients of Saving Rate for Increasing Sample Sizes



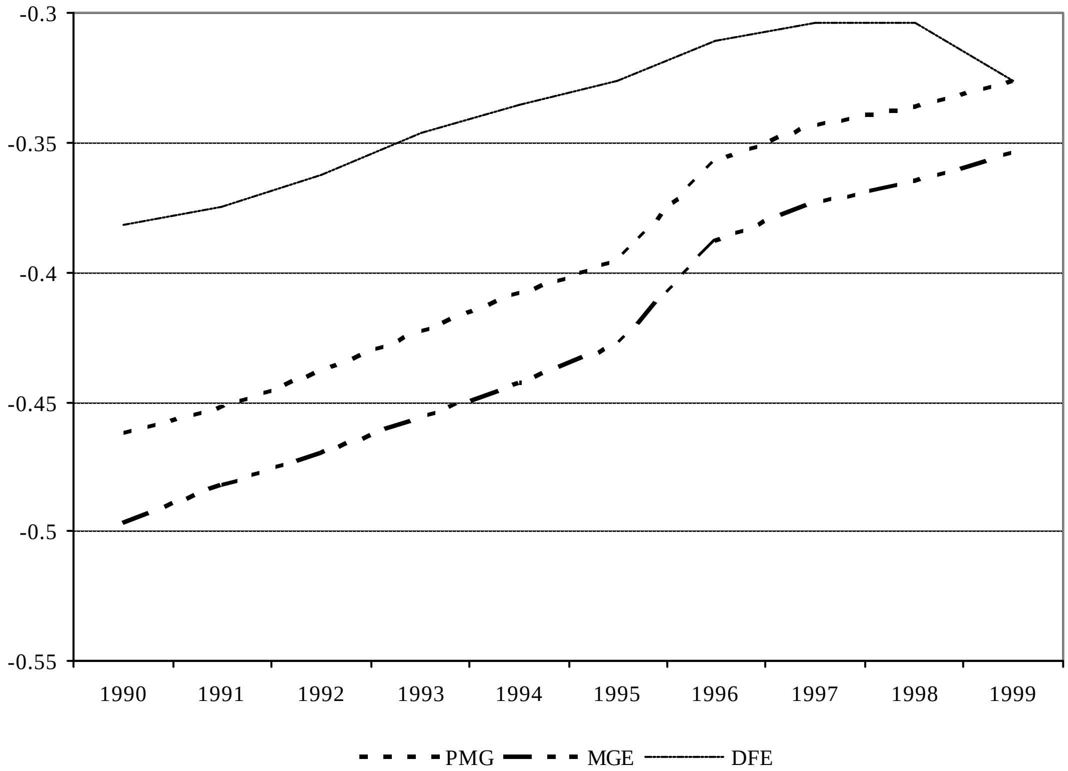
Note: The sample period is extended one year at a time. Thus, the first observation shows the coefficient estimates for the sample from 1960 to 1990 (at 1990 in the chart) and the last one shows the estimates for the sample from 1960 to 1999 (at 1999).

Figure 5: Coefficient of Saving Rate for Moving Samples of Same Size



Note: A moving sample of 30 years is used. Thus, the first observation shows the coefficient estimates for the sample from 1960 to 1990 (at 1990 in the chart) and the last one shows the estimates for the sample from 1970 to 1999 (at 1999). The estimates for the saving rate have a right-hand scale and the ones for the error- correction term have a left-hand scale.

Figure 6: Coefficients of Error-Correction Term for Moving Samples of Same Size



Note: A moving sample of 30 years is used. Thus, the first observation shows the coefficient estimates for the sample from 1960 to 1990 (at 1990 in the chart) and the last one shows the estimates for the sample from 1970 to 1999 (at 1999). The estimates for the saving rate have a right-hand scale and the ones for the error-correction term have a left-hand scale.

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