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NONLINEAR DYNAMICAL METHODS AND TIME SERIES ANALYSIS

Cointegration with Structural Breaks: An Application to the Feldstein-Horioka Puzzle

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Abstract

This paper revisits the well known Feldstein-Horioka saving-investment correlation puzzle from a time series perspective using a sample of 21 OECD countries. We argue that the strong positive correlation between saving and investment as originally identified by Feldstein and Horioka (1980) arises due to the neglect of the nonstationary properties of the variables as well as the failure to account for potential instabilities in the long run relationship between them. Our methodology is based on instability tests recently proposed in Kejriwal and Perron (2006a) as well as the cointegration test in Arai and Kurozumi (2005) extended to allow for multiple breaks under the null hypothesis of cointegration. Our empirical results show that for all countries except Mexico and the U.K., the cointegrating relationship has changed over time; in most cases, the change being towards a lower saving-investment correlation regime. This is perfectly consistent with the recent evidence on international diversification and integration of world capital markets. Finally, we find that while the saving-investment link bears a close relationship with the degree of openness of the country, there seems to be very little evidence in favour of the commonly held view that the correlation varies with the size of the country.

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1 Introduction

Parameter instability in economic models is a common phenomenon. Accordingly, the problem of estimation & inference in models with structural changes has received a great deal of attention in the theoretical econometrics literature. The tools developed have then been applied in numerous empirical studies to detect instability in important macroeconomic & financial relationships. Stock & Watson (1996) provide extensive evidence of widespread parameter instability in a variety of univariate & bivariate macroeconomic time series relations. Bai & Perron (1998) provide an extensive treatment of issues related to estimating & testing multiple structural changes in stationary models.

With nonstationary variables, the case of practical interest is where the variables are individually $I(1)$ but cointegrated. Bai, Lumsdaine & Stock (1998) consider a single break in a multi-equations system & show the estimates obtained by maximizing the likelihood function to be consistent. Kejriwal & Perron (2006b) study the properties of the estimates of the break dates & parameters in a linear regression with multiple structural changes involving $I(1)$, $I(0)$ & trending regressors. With respect to testing, Hansen (1992) develops tests of parameter stability in models where all coefficients are allowed to change. These tests are based on the Lagrange Multiplier (LM) principle.

Kejriwal & Perron (2006a) study issues related to testing for multiple structural changes in cointegrated regression models. They propose a testing procedure that not only enables detection of parameter instability but also allows consistent estimation of the number of breaks. In this paper, the theoretical analysis developed in Kejriwal & Perron (2006a) is applied to analyze the Feldstein-Horioka puzzle which has become quite popular in the international finance literature. In a seminal paper, Feldstein & Horioka (1980) show that across OECD countries, long period averages of national saving rates are highly correlated with similar averages of domestic investment rates. These results were interpreted as evidence of capital immobility within the developed world. However, this interpretation stands in direct contrast with the evidence on integration of world capital markets & liberalization of capital controls.

Figure 1 present a plot of saving & investment rates for a sample of 21 OECD countries. For all the countries, there seems to be a close comovement between the series. However, the plots also suggest that the saving-investment association may have altered over time. Figure 2 present 20 quarter rolling estimates of the coefficient estimate obtained from a regression of the investment rate on the savings rate (& a constant). The dotted lines represent the 95% confidence bands.[1] For most countries, these estimates reveal substantial time variation in the correlation between saving & investment.

We argue that these high saving-investment correlations in Feldstein & Horioka (1980) arise due to the neglect of the nonstationary properties of investment & saving as well as the failure to account for structural breaks in the long run relationship between the variables. Our empirical results show that for all countries except U.K. & Mexico, a regression model which allows for structural changes in the long run relationship between saving & investment rates provides a more adequate specification of the association between the variables than a simple constant parameters model. In particular, we find that for most countries the timing of the breaks as well as the magnitude of the coefficients are in accord with periods of high/low international diversification. Hence, coefficient values in regimes during the '90s are generally smaller than those during the '70s, a period of limited integration of world capital markets.

The paper is organized as follows. Section 2 reviews the Feldstein-Horioka puzzle & the various theoretical & empirical explanations that have been proposed as potential solutions. Section 3 describes the methodology used to test the Feldstein-Horioka hypothesis. Section 4 presents our empirical results. Section 5 contains a discussion of the results. Section 6 offers concluding remarks & all technical material are included in a Technical Appendix.

2 The Feldstein-Horioka Puzzle

Feldstein & Horioka (1980) regressed long averages on the investment-to-output ratio on the saving-to-output ratio, using a cross section of 16 OECD countries over the period 1960-74. They found the following least squares regression result:

$$(\frac{I}{Y})_i = \underset{(0.02)}{0.04} + \underset{(0.07)}{0.89}(\frac{S}{Y})_i \ , \qquad R^2 = 0.91 \tag{1}$$

where $(I/Y)_i$ is the ratio of gross domestic investment to GDP in country i & $(S/Y)_i$ is the corresponding ratio of gross domestic saving to GDP. In order to examine the robustness of the results with respect to the choice of sample period, the regression was also estimated over 1960-64, 1965-69

[1] The confidence bands are constructed using a HAC estimator based on the quadratic spectral kernel and an AR(1) approximation to calculate the bandwidth. (See Andrews, 1991).

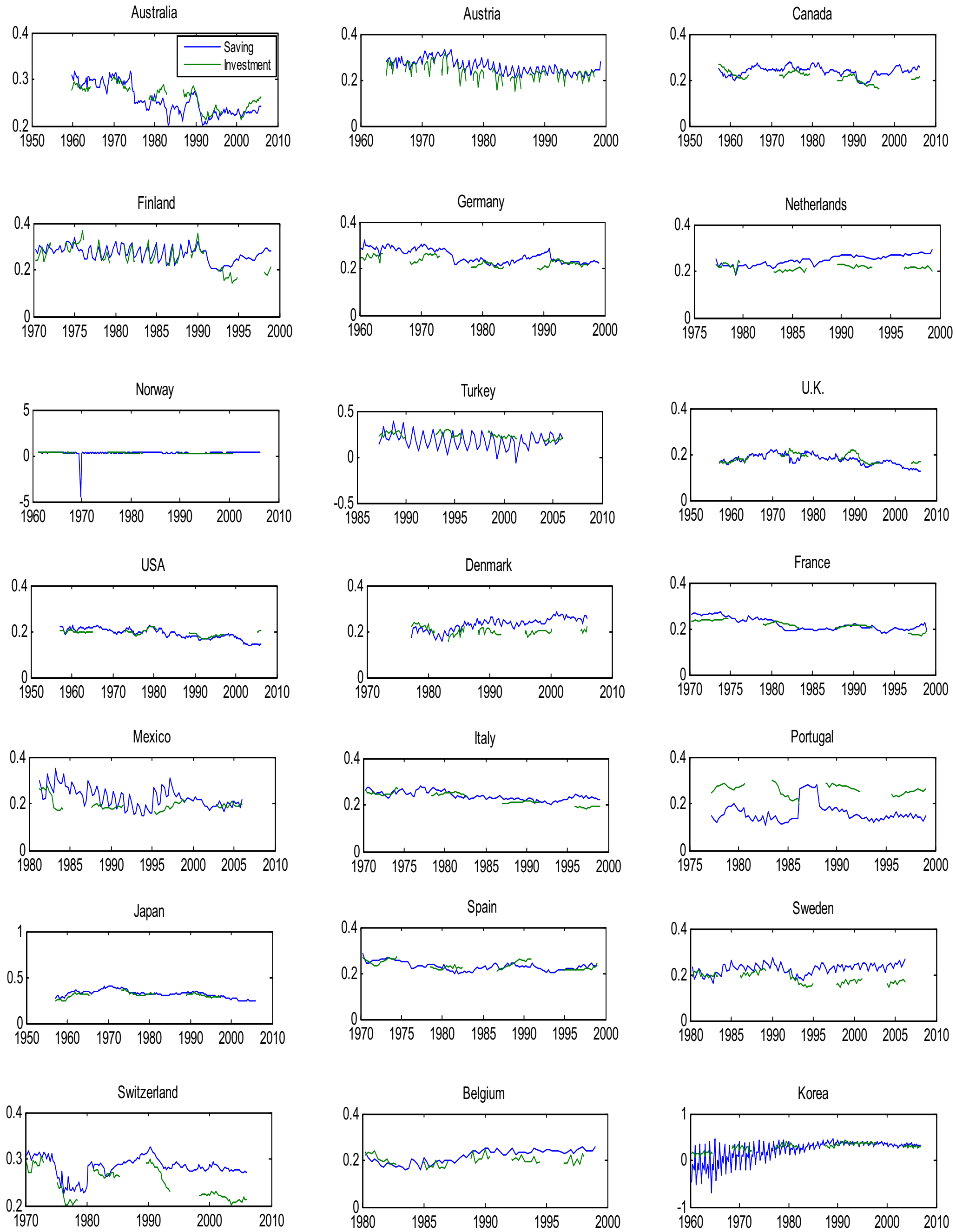

Figure 1: Saving and Investment Rates

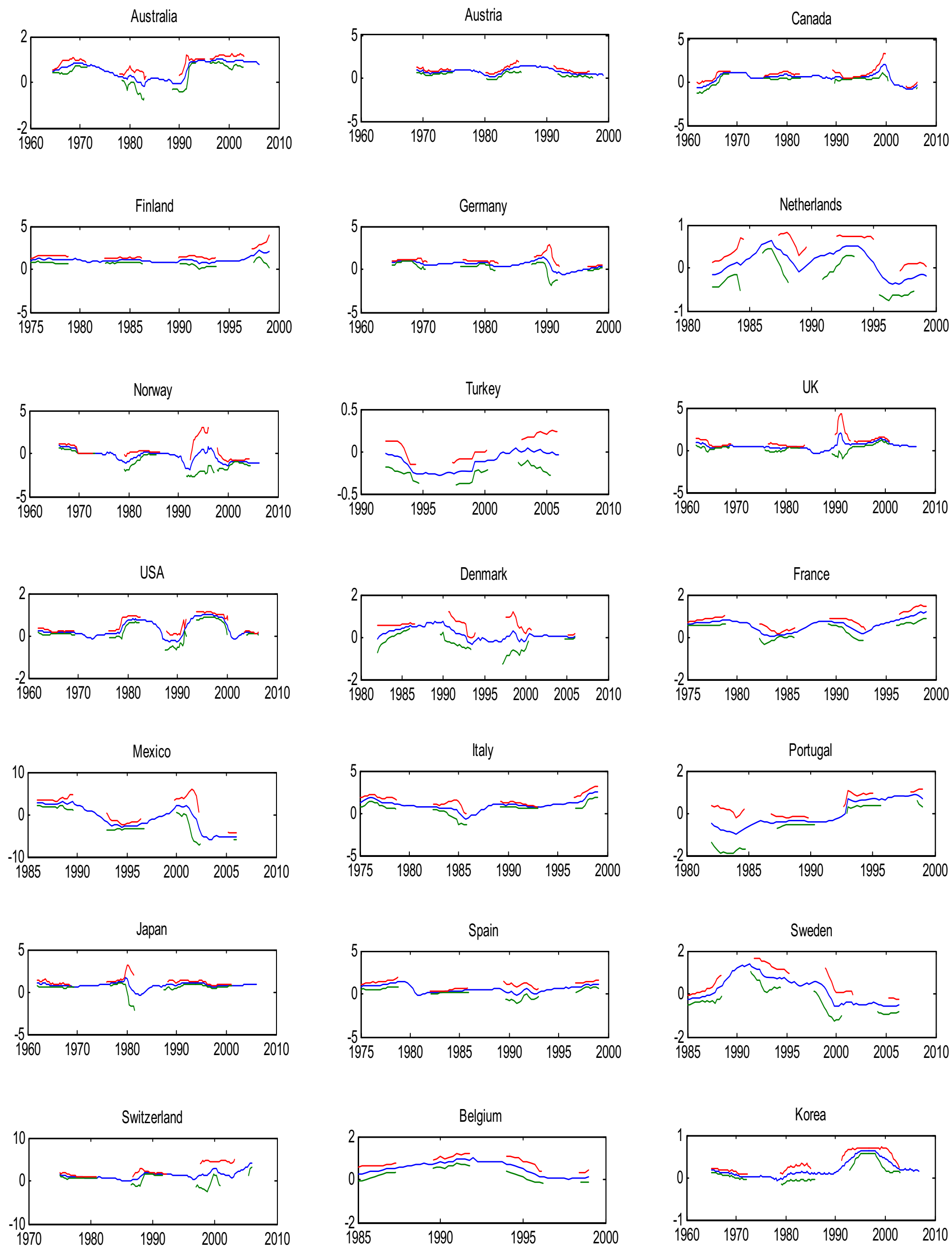

Figure 2: 20 Quarter Rolling Estimates of Regression Coefficients

& 1970-74. Recognizing the potential endogeneity of the saving ratio they also considered a simultaneous equations model estimated by two stage least squares. The parameter estimates were very similar to the OLS estimates & hence they ruled out endogeneity as a possible cause for the high saving-investment association. They also do not find any significant influence of either country size or openness on the saving-investment coefficient. Obstfeld & Rogoff (1996, Chapter 3) estimated the same regression using a sample of 22 countries over the period 1982-91 with the following result:

$$(\frac{I}{Y})_i = \underset{(0.02)}{0.09} + \underset{(0.09)}{0.62}(\frac{S}{Y})_i \ , \qquad R^2 = 0.69 \tag{2}$$

These results were interpreted as evidence of capital immobility for if capital indeed were internationally mobile among industrialized countries, such correlations would be much smaller as a country's savings would then be free to seek out the most productive investment opportunities worldwide. If one accepts this argument, these regression results pose a puzzle, given the evidence on integration of world capital markets & liberalization of capital controls. Recognizing its importance in the literature, Obstfeld & Rogoff (2000a, p.175) call it "the mother of all puzzles". Most explanations that have been suggested tend to be empirically inadequate &, more troublesome still, tend to fix one puzzle at the expense of creating others. In what follows, we will briefly review the various theoretical & empirical explanations that have been suggested as possible solutions to the puzzle.

2.1 Theoretical Studies

A vast body of theoretical work has attempted to show that the observed positive saving-investment correlation does not provide any indication of the degree of capital mobility. Obstfeld (1986) shows that a deterministic dynamic equilibrium model with perfect capital mobility produces positive correlation between saving & investment as a result of persistent productivity changes or population growth. Subsequently, Engel & Kletzer (1989) have argued that the presence of a non-traded consumption good can explain the high correlation even under perfect capital mobility.

In an influential paper, Baxter & Crucini (1993) show that the observed positive association between saving & investment arises naturally within a quantitatively restricted general equilibrium model with perfect mobility of financial & physical capital. Their model is consistent with the observation that saving-investment correlations are larger for larger countries but are still substantial for small countries. The model also predicts that current account deficits are associated with investment booms, thus supporting Sachs' (1981) view that international investment flows are important short run determinants of current account movements. However, the model falls short in two dimensions. First, it predicts that consumption should be nearly perfectly correlated across countries. Second, the one sector international trade model has difficulty generating the positive international comovement of investments, labor inputs & outputs which are observed in the data.

Obstfeld & Rogoff (2000b) propose a potential solution to the Feldstein-Horioka puzzle (as well as other international finance puzzles) by explicitly introducing a friction in the form of a modest & plausible level of international trade costs in goods markets. These trade costs may include transport costs, tariffs, nontariff barriers & possibly other broader factors that impede trade.

2.2 Cross Section & Panel Studies

Since the publication of Feldstein & Horioka's paper, numerous cross section & panel studies have been conducted in order to find a solution to the puzzle. We do not attempt a comprehensive review & focus on the most important contributions. Murphy (1984) studied a cross section of 17 countries & found that saving-investment correlations were larger for larger countries. He found that the average coefficient on the saving-investment coefficient is only 0.59 for the ten smallest countries in his sample, compared with an average coefficient of 0.98 for the seven largest countries. Tesar (1991) showed that the high coefficient is robust to changes in the length of the interval over which the average is taken. Using a sample of 24 OECD countries, she found that for three-year averages the coefficient ranged from 0.76 to 0.95 while for one-year averages the coefficient ranged from 0.67 to 0.97.

Sinn (1992) shows that a standard small open economy model based on an intertemporal approach to the balance of payments implies that the common use of long term averages of saving & investment shares may bias the regression coefficient towards non-rejection of the hypothesis of capital immobility. He finds that coefficients based on annual observations are lower & that they vary considerably from year to year.

In a recent paper, Banerjee & Zanghieri (2004) apply the recently developed theory of panel unit roots & cointegration to study the saving-investment association in a panel of 14 European countries. They consider subsets of countries in the panel as well as apply tests to various disaggregates & transformations of the core datasets. They also interpret their findings in the light of the presence of cross country cointegration. Specifically, they argue that cross country cointegration is likely to play a critical role in the finding that, when panel unit root tests are employed, the current account turns out to be stationary.

2.3 Time Series Studies

In the working paper version of their paper, Feldstein & Horioka computed annual times series estimates of the coefficient for each of 21 countries. They averaged 0.64 but showed great variation across countries. They argue that the simultaneity bias makes these time series estimates too unreliable. Nevertheless, economists have continued to analyze the problem in a time series framework. Obstfeld (1986) computes time series correlations between changes in saving & investment rates using quarterly data from 7 OECD countries & finds correlations ranging from 0.13 to 0.91. Tesar (1991) provides time series plots of the saving & investment ratios over the full sample period. Within each country, it appears that saving & investment are highly correlated at the annual frequency. Frankel (1991) finds evidence that the U.S. saving-investment coefficient fell in the 1980's. Frankel (1992) points out that if the saving-investment regressions were a good test for barriers to financial-market integration, one would expect to see the regression coefficient falling over time.

3 Methodology

In contrast to Baxter & Crucini (1993) who try to provide a theoretical justification of the observed saving-investment association, we take the empirical route in that we argue that these observed correlations may be overstated due to neglecting the nonstationary properties of the saving & investment rates as well as due to the failure to take into account possible structural changes that may have affected the long run relationship over the period under consideration. Our choice of studying the time series correlation between saving & investment for individual countries rather than the cross section association is motivated by several considerations. First, using long-term averages of saving & investment rates may suggest a long run relationship even when no correlation exists (Sinn, 1992). Second, the cross section analysis may be subject to sample selection bias.[2] Third, in a cross section approach, the growth rate

[2]For example, in the study of Tesar (1991) when Luxembourg is excluded from the sample, the findings change dramatically, the correlation between saving and investment increasing from 0.35 to 0.84.

of income or the presence of a non-traded consumption good may simultaneously affect saving & investment. Feldstein & Horioka (1980) & Dooley et al. (1987) both address the potential simultaneity of saving & investment shares by instrumental variable analysis. An advantage of the cointegration setup used in this paper is that the potential endogeneity of the regressors can be handled simply by augmenting the cointegrating regression with leads & lags of the first differences of the endogenous regressor (here the savings rate). This allows standard asymptotic inference about the regression coefficients. Of course it is possible that using saving as the only regressor may result in omission of certain relevant regressors which may bias the cointegration tests towards rejection of the hypothesis of cointegration between saving & investment. However, our empirical analysis finds evidence of cointegration for all countries except U.K.. Under cointegration, the estimate of the cointegrating vector is consistent even in the presence of endogeneity. It is also possible that there are other cointegrating vectors which involve these omitted variables. However, we do not consider these other vectors since our main focus is on analyzing the long run saving-investment relationship.

The methodology used in this paper is to first use unit root tests to verify that the saving & investment rates are individually integrated of order one. It may be argued that since the rates are bounded by 0 & 1 they may be persistent rather than $I(1)$ processes. It is, however, a common practice to model such bounded persistent series as $I(1)$ rather than stationary. For example, though the nominal interest rate is bounded, a wide range of macroeconomic studies model the process as possessing a unit root. Nicolau (2002) points out that while it is not possible to say that these bounded time series are random walks because random walks are limitless with probability one (as time goes to infinity), some of these time series behave just like random walks. He shows that the paths of such bounded random walks are almost indistinguishable from usual random walks, although they are stochastically bounded by an upper & lower finite limit. Cavaliere (2005) develops an asymptotic theory for integrated & near-integrated time series whose range is constrained in some ways. He introduces the bounded unit root distribution to describe the limiting distribution of the sample first-order autoregressive coefficient of a random walk under range constraints. His theoretical results show that the presence of such constraints can lead to drastically different asymptotics. We simulated the critical values corresponding to his bounded unit root distribution for the countries considered & found that the critical values are the same as that of the standard unit root tests. Hence, the 0-1 bounds on the saving & investment shares are not constraining in any way.

We then test the stability of the saving-investment relationship using the tests proposed in Kejriwal & Perron (2006a). Given that these tests can reject the null of stability when the regression is really a spurious one, we need to verify that the variables are indeed cointegrated. In this sense, these cointegration tests are used as confirmatory tests. If we find evidence of cointegration, we select the number of breaks using both the sequential procedure outlined in Kejriwal & Perron (2006a) as well as information criteria. Hence, tests for breaks in the long run relationship should be used in conjunction with tests for the presence/absence of cointegration allowing for structural changes in the coefficients. Finally, we estimate the model incorporating the breaks in order to study how the saving-investment association may have altered over time.

3.1 Unit Root Tests

Perron & Ng (1996) analyze a class of modified unit root tests & show that these tests are far more robust to size distortions than other unit root tests in the literature, especially when the residuals have negative serial correlation. Ng & Perron (2001) apply the idea of GLS detrending to the modified tests & show that accurate size & non-negligible power gains can be obtained when used in conjunction with an autoregressive spectral density estimator at frequency zero provided the truncation lag is appropriately selected. A detailed description of the various tests can be found in the Technical Appendix.

One problem with the above class of tests is that for non-local alternatives the power can be very small. To alleviate this problem, Perron & Qu (2007) suggest an easy solution which also leads to tests having an exact size even closer, in most cases, to nominal size. It involves using OLS instead of GLS detrended data when constructing the modified information criteria.

It is by now well known that structural changes in the mean of a stationary time series biases the usual tests for a unit root towards non-rejection;.see, for example, Perron (1990). Hansen (2000) argues that the LM test is quite poorly behaved in the presence of structural changes in the marginal distribution of the regressors. However, the sup-Wald test is shown to be reasonably robust to such shifts. Since the tests proposed in Kejriwal & Perron (2006a) are modified versions of the sup Wald test, such shifts are unlikely to have any significant effect on the finite sample properties of the tests.

3.2 Structural Break Tests

Having verified the presence of unit roots, the next step entails testing whether the saving-investment relationship has remained stable through time. Hansen (1992) develops sup and mean LM tests of the null hypothesis of no change in cointegrated models where all coefficients are allowed to change. These tests are directed against the alternative of a one time change in parameters.

A potential problem with the application of these LM type tests is that they exhibit non-monotonic power in finite samples. This means that as the magnitude of change under the alternative hypothesis increases, power can decline sharply so that it may not be possible to detect large changes using this class of tests. This problem arises due to the estimation of the long run variance of the errors under the null hypothesis of stability. An increase in the break magnitude leads to an increase in the bandwidth which in turn increases the long run variance estimate thus reducing the power of the LM tests. In fact, for certain configurations of parameter changes, power can go to zero as the magnitude of change increases. Moreover, such non-monotonic power functions can also arise when the DGP involves more than one break. An example is when the DGP changes in such a way that the first & third regimes are identical. It is thus useful to develop tests for multiple structural changes.

Kejriwal & Perron (2006a) study issues related to structural changes in cointegrated regression models allowing for both $I(1)$ and $I(0)$ regressors. They derive the limit distribution of the sup Wald test for the null hypothesis of no structural change in a general model which allows both $I(1)$ & $I(0)$ regressors as well as multiple breaks. They also propose a sequential procedure which permits consistent estimation of the number of breaks. For simplicity, suppose that the model under consideration involves only $I(1)$ regressors (as in our case). Let SSR_0 denote the sum of squared residuals under the null hypothesis of no breaks & SSR_k denote the sum of squared residuals under the alternative hypothesis of k breaks. We denote $\lambda = \{\lambda_1, ..., \lambda_m\}$ as the vector of break fractions defined by $\lambda_i = T_i/T$ for $i = 1, ..., m$, T_i being the corresponding break dates. The tests proposed are as follows:

$$
\begin{aligned}
\sup\ F_T^*(k) &= \sup_{\lambda \in \Lambda_\epsilon} \frac{SSR_0 - SSR_k}{\hat{\sigma}^2} \\
UD\max F_T^*(M) &= \max_{1 \leq k \leq m}\ F_T^*(k)
\end{aligned}
$$

where

$$\hat{\sigma}^2 = T^{-1}\sum_{t=1}^{T}\tilde{u}_t^2 + 2T^{-1}\sum_{j=1}^{T-1} w(j/\hat{h})\sum_{t=j+1}^{T}\tilde{u}_t\tilde{u}_{t-j} \tag{3}$$

& $\tilde{u}_t$ $(t = 1, ..., T)$ are the residuals from the model estimated under the null hypothesis of no structural change. Also, for some arbitrary small positive number ϵ, $\Lambda_\epsilon = \{\lambda : |\lambda_{i+1} - \lambda_i| \geq \epsilon, \lambda_1 \geq \epsilon, \lambda_k \leq 1 - \epsilon\}$. Using Andrews' (1991) automatic bandwidth estimator with a quadratic spectral kernel, we have $\hat{h} = 1.3221(\hat{a}(2)T)^{1/5}$, $\hat{a}(2) = 4\hat{\rho}^2/(1-\hat{\rho})^4$ & $\hat{\rho} = \sum_{t=2}^{T}\hat{u}_t\hat{u}_{t-1}/\sum_{t=2}^{T}\hat{u}_{t-1}^2$, where $\hat{u}_t$ $(t = 1, ..., T)$ are the residuals from the model estimated under the alternative hypothesis. Kejriwal & Perron (2006a) conduct simulation experiments to show that the power functions of the suggested test are monotonic while the size remains adequate. This justifies the use of the proposed test as opposed to the LM type tests in empirical applications.

One criticism of the time series approach, as pointed out by Feldstein & Horioka, is that the resulting simultaneity bias may be quite serious so as to render the coefficient estimates unreliable. With cointegration, the parameter estimates are consistent even in the presence of endogeneity but not optimal. In order to circumvent this problem of endogeneity, we use the dynamic OLS estimator which involves augmenting the OLS regression with leads & lags of the first differences of the regressor (here the saving rate).[3] The number of leads & lags is set equal to 2. Kejriwal & Perron (2006a, 2007) show that the lead-lag length can also be selected using data dependent rules such as information criteria.

In addition to the tests above, Kejriwal & Perron (2006a) consider a test of the null hypothesis of k breaks against the alternative that an additional break exists. In fact, the test is equivalent to the application of $(k+1)$ tests of the null hypothesis of no structural change versus the alternative hypothesis of a single change. More specifically, it is defined by

$$SEQ_T(k+1|k) = \max_{1\ \leq\ j\ \leq\ k+1}\ \sup_{\tau\in\Lambda_{j,\epsilon}}\left\{\frac{A_T(k) - B_T(\tau,k)}{\hat{\sigma}_{k+1}^2}\right\}$$

where $A_T(k) = SSR_T(\hat{T}_1, ..., \hat{T}_k)$, $B_T(\tau,k) = SSR_T(\hat{T}_1, ..., \hat{T}_{j-1}, \tau, \hat{T}_j, ..., \hat{T}_k)$, $\Lambda_{j,\epsilon} = \left\{\tau : \hat{T}_{j-1} + (\hat{T}_j - \hat{T}_{j-1})\epsilon \leq \tau \leq \hat{T}_j - (\hat{T}_j - \hat{T}_{j-1})\epsilon\right\}$ & $\hat{\sigma}_{k+1}^2$ is a consistent estimator of the long run variance under the null hypothesis but where the bandwidth is estimated using residuals under the alternative hypothesis as

[3] See Saikkonen (1991), Stock and Watson (1993).

in (3). The procedure is implemented as follows: First we test for zero versus one break; if a rejection occurs we test for one versus two breaks & so on until a non-rejection occurs. The number of breaks is estimated as the number of rejections. Such a sequential testing procedure provides a consistent estimate of the true number of breaks. (Bai & Perron, 1998).

Bai & Perron (2006) conducted simulation experiments to show that the one break test may have low power against the alternative hypothesis of two breaks when the parameter values change in such a way that the first & third regimes are identical. In such a situation, the sequential procedure might end up selecting no breaks. Hence, a useful strategy is to first see whether the $UD\max$ test is significant or not. If it is, we can then use the sequential procedure to choose the number of breaks. This is the strategy we adopt in this paper.

As an alternative to the sequential procedure, the number of breaks may also be selected using information criteria. Yao (1988) suggests the use of the Bayesian Information Criterion (BIC) defined as

$$BIC(m) = \ln \hat{\sigma}^2(m) + p^* \ln(T)/T,$$

where $p^* = (m+1)q+m+p$, & $\hat{\sigma}^2(m) = T^{-1}S_T(\hat{T}_1, ..., \hat{T}_m)$, $\hat{T}_1, ..., \hat{T}_m$ denoting the estimated break dates & $S_T(\hat{T}_1, ..., \hat{T}_m)$ the sum of squared residuals under m breaks. Also, q is the number of coefficients which are allowed to change & p is the number of coefficients that are held fixed. Liu, Wu & Zidek (1997) propose a modified Schwarz' Criterion that takes the form

$$LWZ(m) = \ln(S_T(\hat{T}_1, ..., \hat{T}_m)/(T - p^*)) + (p^*/T)c_0(\ln(T))^{2+\delta_0}$$

They suggest using $\delta_0 = 0.1$ & $c_0 = 0.299$. We do not consider estimating the number of breaks using the Akaike's Information Criterion (AIC) because it has been shown to perform quite poorly in the presence of serial correlation (see Perron, 1997). In this paper, we will use the sequential procedure as well as the information criteria in order to detect the number of breaks in the saving-investment relationship.

3.3 Cointegration Tests

Since the structural change tests also have power against a purely spurious regression, we need to verify that the variables are indeed cointegrated. Phillips & Ouliaris (1990) develop an asymptotic theory for residual based tests of the null hypothesis of no cointegration. Gregory & Hansen (1996) & Gregory, Nason & Watt (1996) show through Monte Carlo experiments that the power of the conventional ADF test falls sharply when there is a break in the cointegrating relationship. The idea is essentially the same as that pointed out by Perron (1989) in the context of unit root tests. In particular, the presence of unaccounted shifts in the long run relationship biases the usual cointegration tests in favour of non-rejection of the null hypothesis of no cointegration. Hence, Gregory & Hansen (1996) examine tests for cointegration which allow for the possibility of regime shifts. In particular, they propose ADF, Z_α & Z_t-type tests designed to test the null of no cointegration against the alternative of cointegration in the presence of a possible regime shift. In particular, they consider cases where the intercept &/or slope coefficients have a single break of unknown timing. The test statistics ADF^*, Z_α^* & Z_t^* are computed as the minimal values of the usual statistics over all possible breakpoints. For countries where at least one of the three procedures to select the number of breaks choose a single break, we present results for the Gregory-Hansen (henceforth G-H) tests. A rejection by these tests would then confirm the presence of a cointegrating relationship. However, as in the case of unit root tests, the value of the break date associated with the minimal value of a given statistic is not, in general, a consistent estimate of the break date if a change is present. Moreover, these tests are designed to have power against the alternative of a single break in parameters & hence may have low power when the alternative involves more than one break. Finally, from the view of classical hypothesis testing, if we are primarily concerned about cointegration with structural breaks, cointegration seems a more natural choice for the null hypothesis.

To avoid the problems encountered in the application of the G-H tests, we consider the residual based test of the null hypothesis of cointegration with structural breaks proposed in Arai & Kurozumi (2005). They propose the LM test based on partial sums of residuals where the break point is obtained by minimizing the sum of squared residuals. The limiting properties of the break point ensure that the test statistic has the same distribution as the known break case. However, the test is restrictive in the sense that only a single structural break is considered under the null hypothesis. Thus, the test may tend to reject the null of cointegration when the true data generating process exhibits cointegration with multiple breaks. Hence, we extend their test by incorporating multiple breaks under the null hypothesis. Specifically, we derive the limit distribution & simulate critical values for those combinations of break dates which are relevant for our empirical application. Arai & Kurozumi (2005) consider 3 models: (a) Level Shift, (b) Level Shift with Trend & (c) Regime Shift. Since theory suggests that the variables in question do not contain a trend & what we are really interested in is the stability of the saving-investment coefficient, we consider only model (c), which we may write as follows:

$$y_t = c_i + z_t'\beta_i + u_t \qquad if\ T_{i-1} < t \leq T_i \tag{4}$$

for $i = 1, ..., k+1$, where k is the number of breaks, z_t is a q vector of $I(1)$ regressors & where, by convention, $T_0 = 0$ & $T_{k+1} = T$. In order to correct for the potential endogeneity of the regressors, we augment (4) with the leads & lags of the first differences of the $I(1)$ regressors. The augmented regression is written as

$$y_t = c_i + z_t'\beta_i + \sum_{j=-l_T}^{l_T} \Delta z_{t-j}'\Pi_j + u_t^* \qquad if\ T_{i-1} < t \leq T_i \tag{5}$$

The test statistic is given by

$$\widetilde{V}_1(\hat{\lambda}) = (T^{-2}\sum_{t=1}^{T} S_t(\hat{\lambda})^2)/\hat{\Omega}_{11}$$

where $\hat{\Omega}_{11}$ is a consistent estimate of the long run variance of u_t^*, $\hat{\lambda} = (\hat{T}_1/T, ..., \hat{T}_k/T)$ & $(\hat{T}_1, ..., \hat{T}_k)$ are obtained by minimizing the sum of squared residuals. The break dates are obtained using the dynamic programming algorithm proposed in Bai & Perron (2003). We denote the test statistic with k breaks as $\widetilde{V}_k(\hat{\lambda})$.

Let $\xi_t = (u_t,\ \eta_t')'$ where $\Delta z_t = \eta_t$. We assume ξ_t satisfies a multivariate functional central limit theorem (FCLT); see Arai & Kurozumi (2005) & Kejriwal & Perron (2006a) for details. The limit distribution of the test under the null hypothesis is stated in the following proposition proved in the Technical Appendix.

Proposition *Assume that the data are generated by (4) & the z_t's are strictly exogenous. Then under the null hypothesis, we have*

$$\widetilde{V}_k(\hat{\lambda}) \Rightarrow \int_0^1 Q_{\lambda,k}^2(r)dr$$

uniformly over $\lambda \in [0,1]^k$ where $Q_{\lambda,k}(r) = W_1(r) - G(r)$ where $G(r) = \int_0^r W(s)' \left\{ \sum_{i=1}^{k+1} \left(\int_{\lambda_{i-1}}^{\lambda_i} WW' \right)^{-1} \left(\int_{\lambda_{i-1}}^{\lambda_i} WdW_1 \right) I(\lambda_{i-1} < s \leq \lambda_i) \right\} ds$ & $W(s) = (1, W_2(r)')'$, W_2 being a vector of q independent Wiener processes & W_1 is a one dimensional Wiener process independent of W_2.

Arai & Kurozumi (2005) show, in the single break case, that the limit distribution of the test statistic based on the estimated break date is the same as that when the break date is known. This is because the rate of convergence of the estimated break fraction is fast enough so that asymptotically it does not matter whether we use the estimated or true break date when constructing the test statistic. Critical values are obtained by simulation using 500 steps & 2000 replications. The Wiener processes are approximated by partial sums of *i.i.d.* $N(0, 1)$ random variables. Note that critical values will be different depending on the particular combination of break fractions. For the purpose of our empirical application, we will simulate critical values for those combinations that are obtained by minimizing the sum of squared residuals for the countries in question.

A caveat associated with this class of tests, as pointed out by Perron (2006), is that if cointegration actually holds, there must be a change in the cointegrating relationship. This is because the search for the potential break date is restricted to break fractions that are bounded, in large samples, from the boundaries 0 & 1. Hence, we use the instability tests proposed in Kejriwal & Perron (2006a) as well as information criteria to ensure the existence of breaks.

4 Empirical Results

We use quarterly data obtained from the IMF's *International Financial Statistics*. The sample period varies across countries depending on the availability of the data.[4] For our measure of investment, we use gross fixed capital formation. Bayoumi (1990) points out the advantage of gross fixed capital formation as a measure of investment - it has a lesser tendency to behave procyclically because it excludes the highly procyclical inventories component. As a measure of saving, we use what Baxter & Crucini (1993) refer to as "basic saving". Basic saving is defined as GDP less two types of consumption - private & government consumption. As previously noted by Obstfeld (1986) & Stockman & Svensson (1987), the national income accounts (NIA) measure of saving can differ from true saving. The difference arises when foreigners

[4]The sample periods for the different countries are as follows- Australia-1959:Q3-2005:Q4; Austria-1964:Q1-1999:Q1; Canada, U.K. & U.S.A.-1957:Q1-2006:Q1; Finland, France, Italy, Spain-1970:Q1-1998:Q4; Germany-1960:Q1-1998:Q4; Netherlands-1977:Q1-1999:Q1; Norway-1961:Q1-2006Q1; Turkey-1987:Q1-2005:Q4; Denmark-1977:Q1-2005:Q4; Mexico-1981:Q1-2005:Q4; Portugal- 1977:Q1-1998:Q4; Japan-1957:Q1-2005:Q4; Sweden-1980:Q1-2006:Q1; Switzerland-1970:Q1-2005:Q4; Belgium-1980:Q1-1998:Q4; Korea-1960:Q1-2006:Q3.

own shares in domestic firms & when firms finance expenditure from retained earnings. The advantage of using basic saving is that it is invariant to different assumptions about firms' financing decisions. Both saving & investment are expressed as percentages of GDP.

Before we begin to implement our methodology, it is useful to have an idea of the results one would obtain by simply running an OLS regression of the investment rate on the saving rate. The difference between these results & those obtained using our proposed methodology will allow us to highlight the importance of a correct empirical specification when analyzing the relation between saving & investment rates. To that end, Table 1 presents coefficient estimates as well as the R^2 from a regression of the investment rate on the saving rate.[5] The estimates in all countries except Netherlands, Norway, Turkey, Denmark, Mexico, Portugal & Sweden are significant at the 1% level.[6] Hence, on the basis of such an analysis, one might be tempted to conclude in favour of a close association between the two variables.

Table 2 present the results of unit root tests for the individual saving & investment series. The saving & investment series plotted in Figure 1 do not seem to possess a deterministic time trend. Moreover, macroeconomic theory suggests that these ratios do not have a trend. Hence, the unit root tests are conducted with a constant as the only deterministic component in the regression. For Australia, Austria, Germany, Netherlands, U.K., U.S.A., France, Mexico, Italy, Japan, Spain, Belgium & Korea, none of the unit root tests reject the null hypothesis of a unit root in saving & investment. For Canada, Finland, Norway & Switzerland while investment clearly has a unit root, the null of a unit root in saving is rejected. For Turkey, Denmark, Portugal & Sweden a unit root in investment is rejected at the 5% level but not at the 1% level.

Having verified the nonstationarity of saving & investment rates, we now consider the tests for structural change that have been proposed in Kejriwal & Perron (2006a). We use 15% trimming so that the maximum number of breaks allowed under the alternative hypothesis is 5. Both the intercept & the slope are allowed to change. Table 3 presents results of stability tests as well as number of breaks selected by the sequential procedure & the information criteria BIC and LWZ (The procedures are denoted by (S), (B), (L) respectively). Given the span of data, it seems unreasonable to expect the occurrence of four or more breaks. Hence, as a rule of thumb, these cases are treated as evidence

[5] In all Tables, #, *, ** denote significance at the 10%, 5% & 1% level respectively.

[6] To evaluate significance, we use a 2-sided t test. The correction for potential serial correlation is made using a HAC estimator based on the quadratic spectral kernel and an AR(1) approximation to calculate the bandwidth.

Table 1: OLS Regression of I/Y on S/Y: Estimated Coefficients & Standard Errors

(A)	Aus	Austria	Can	Fin	Ger	Net	Nor	Tur	U.K.	U.S.A.
Estimate	0.64**	0.79**	0.47**	1.04**	0.55**	-0.01	0.00	-0.09*	0.46**	0.24**
Standard Error	0.003	0.091	0.165	0.170	0.114	0.087	0.005	0.045	0.042	0.047
R^2	0.69	0.38	0.21	0.44	0.47	0.00	0.00	0.04	0.38	0.26

(B)	Den	Fra	Mex	Ita	Por	Jap	Spa	Swe	Swi	Bel	Kor
Estimate	0.11	0.63**	0.01	1.03**	-0.30*	0.82**	0.67**	0.09	1.05**	0.40**	0.16**
Standard Error	0.121	0.089	0.102	0.149	0.131	0.047	0.152	0.181	0.143	0.084	0.020
R^2	0.02	0.63	0.00	0.48	0.16	0.88	0.32	0.01	0.53	0.25	0.23

Table 2: Unit Root Tests

(A)	Series	Aus	Austria	Can	Fin	Ger	Net	Nor	Tur	U.K.	U.S.A.
MZ_α^{GLS}	I/Y	-5.26	-1.88	-0.61	-4.62	-5.93	-5.14	-0.83	-6.21*	-4.21	-4.09
	S/Y	-0.34	-3.11	-10.72*	-11.96*	-1.59	-0.77	-10.71*	-1.23	-3.87	-0.99
MSB^{GLS}	I/Y	0.31	0.45	0.70	0.32	0.28	0.28	0.46	0.28	0.34	0.35
	S/Y	0.74	0.40	0.21*	.20*	0.44	0.46	0.22*	0.62	0.31	0.40
MZ_t^{GLS}	I/Y	-1.61	-0.84	-0.43	-1.49	-1.64	-1.45	-0.38	-1.75	-1.45	-1.43
	S/Y	-0.25	-1.24	-2.31*	-2.45*	-0.69	-0.35	-2.31*	-0.76	-1.21	-0.40
MP_T^{GLS}	I/Y	4.70	11.50	-26.86	5.38	4.41	5.18	14.89	3.99	5.82	5.99
	S/Y	31.56	7.89	2.31*	2.05*	12.10	14.94	2.30*	19.07	6.50	12.64

Table 2: Unit Root Tests (Contd.)

(B)	Series	Den	Fra	Mex	Ita	Por	Jap	Spa	Swe	Swi	Bel	Kor
MZ_{α}^{GLS}	I/Y	-5.86*	-1.64	-2.35	0.52	-6.86*	-2.60	-3.76	-6.18*	-3.39	-3.15	-0.42
	S/Y	-0.14	-0.18	-1.40	-2.37	-8.99**	-1.65	-0.78	-8.06*	-8.44**	-1.21	0.54
MSB^{GLS}	I/Y	0.29*	0.44	0.42	0.81	0.27*	0.42	0.36	0.28	0.33	0.39	0.79
	S/Y	0.70	0.61	0.74	0.39	0.24**	0.52	0.67	-1.85*	0.23*	0.56	3.21
MZ_{t}^{GLS}	I/Y	-1.70*	-0.72	-0.99	0.43	-1.84*	-1.12	-1.36	-1.70*	-1.12	-1.22	-0.34
	S/Y	-0.09	-0.11	0.53	-0.92	-2.12**	-0.86	-0.52	0.23**	-1.99**	-0.69	1.75
MP_{T}^{GLS}	I/Y	4.21*	12.10	9.85	44.45	3.61*	9.32	6.53	4.14*	7.17	7.73	33.95
	S/Y	29.98	24.34	15.19	9.37	2.73**	13.91	24.13	3.63*	3.15**	17.25	591.98

Table 3: Structural Break Tests

(A)	Aus	Austria	Can	Fin	Ger	Net	Nor	Tur	U.K.	U.S.A.
Sup $F^*(1)$	4.87	17.95**	6.37	28.20**	7.08	4.11	15.51*	27.87**	3.01	15.38*
Sup $F^*(2)$	4.06	21.37**	3.64	18.00**	4.84	5.51	10.53*	11.77*	3.37	3.50
Sup $F^*(3)$	3.80	13.88**	2.90	16.55**	3.47	7.36	9.96*	12.19**	2.96	4.34
Sup $F^*(4)$	3.27	10.38**	2.35	12.49**	2.88	10.45**	10.09**	12.39**	6.51	3.55
Sup $F^*(5)$	2.67	8.58**	2.21	17.11**	2.70	10.57**	8.94**	8.85**	5.19	2.79
UDmax	4.87	21.37**	6.37	28.20**	7.08	$10.57^{\#}$	15.51*	27.87**	6.51	15.38*
(S)	0	2	0	1	0	1	1	1	0	1
(B)	4	2	5	2	5	4	4	1	4	3
(L)	3	2	1	1	2	4	3	1	4	3

Table 3: Structural Break Tests (Contd.)

(B)	Den	Fra	Mex	Ita	Por	Jap	Spa	Swe	Swi	Bel	Kor
Sup $F^*(1)$	7.92	7.54	2.44	7.70	5.74	4.93	2.31	21.69**	17.34**	15.68*	8.19
Sup $F^*(2)$	8.68	5.81	2.65	4.44	8.36	4.86	3.16	41.76**	9.94#	11.73*	7.62
Sup $F^*(3)$	10.14*	5.31	2.53	3.72	5.93	3.92	4.02	19.72**	7.45	8.31#	10.55#
Sup $F^*(4)$	8.86**	4.31	2.62	2.95	4.40	3.57	3.25	26.70**	6.40	6.53	9.66
Sup $F^*(5)$	8.01**	2.98	2.32	2.41	3.42	2.90	2.34	10.04**	6.15	5.40#	5.88
UDmax	10.14	7.54	2.65	7.70	8.36	4.93	4.02	41.76**	17.34*	15.68#	10.55#
(S)	0	0	0	0	0	0	0	2	1	2	1
(B)	4	4	4	3	3	4	3	2	4	2	4
(L)	2	3	0	3	2	2	3	2	1	0	3

Table 4: PP and ADF Cointegration Tests

(A)	Aus	Can	Ger	U.K.
Z_t	-4.20**	-2.96	-3.13	-2.97
Z_α	-30.96**	-11.39	-17.71	-16.23
ADF_t	-2.04	-3.00	-2.79	-2.61

(B)	Den	Fra	Mex	Ita	Por	Jap	Spa	Bel
Z_t	-4.33**	-2.40	-3.67*	-2.95	-1.80	-3.78*	-3.08	-6.69**
Z_α	-27.30*	-10.72	-17.78*	-16.13	-6.05	-25.77*	-17.34	-53.51**
ADF_t	-3.29#	-2.40	-4.05**	-1.74	-2.19	-3.15#	-2.43	-4.46**

Table 5: Cointegration Tests with a Single Break

(A): G-H	Can	Fin	Net	Nor	Tur	U.S.A.	Swi	Kor
Z_t^*	-3.71	-7.72**	-5.48**	-6.90**	-7.58**	-2.50	-8.70**	-10.31**
Z_α^*	-20.76	-77.76**	-40.57	-69.64**	-65.79**	-13.11	-97.11**	-128.20**
ADF_t^*	-5.05*	-3.87	-4.77$^{\#}$	-4.21	-4.54	-3.76	-3.66	-3.91

(B): A-K	Can	Fin	Net	Nor	Tur	U.S.A.	Swi	Kor
$V_1(\hat{\lambda})$	0.08	0.13	0.15	0.32**	0.07	0.09	0.17*	0.14$^{\#}$
Break Fraction	0.51	0.80	0.17	0.64	0.75	0.40	0.63	0.64
Break Date	82:Q1	92:Q4	81:Q1	89:Q4	00:Q4	76:Q4	92:Q3	89:Q4

Table 6: Arai-Kurozumi Cointegration Tests with Multiple Breaks

(A)	$\tilde{V}_2(\hat{\lambda})$	$\hat{\lambda}_1$	$\hat{\lambda}_2$	(B)	$\tilde{V}_3(\hat{\lambda})$	$\hat{\lambda}_1$	$\hat{\lambda}_2$	$\hat{\lambda}_3$
Austria	0.04	0.33	0.68	Aus	0.07*	0.28	0.45	0.61
Fin	0.19**	0.26	0.80	Nor	0.13	0.28	0.43	0.64
Ger	0.05	0.34	0.80	U.S.A.	0.06	0.41	0.65	0.85
Den	0.09	0.19	0.34	Fra	0.05	0.25	0.45	0.85
Por	0.11*	0.30	0.49	Ita	0.06$^{\#}$	0.19	0.44	0.80
Jap	0.06	0.32	0.53	Por	0.05	0.30	0.49	0.67
Swe	0.10$^{\#}$	0.30	0.49	Spa	0.05	0.29	0.60	0.75
Belgium	0.05	0.15	0.72	Korea	0.06	0.46	0.65	0.81

in favor of a spurious regression rather than cointegration with breaks (see Kejriwal & Perron, 2006a). For Australia, none of the tests are significant indicating a stable cointegrating relationship. The sequential procedure selects no break. For Austria, the sequential procedure based on the modified tests selects two breaks, a result supported by the information criteria. The test results for Canada do not suggest any instability although the information criterion LWZ selects one break. In the case of Finland, all the modified tests reject stability at the 1% level. The sequential procedure selects a single break. For Germany, none of the tests provide any evidence of instability although LWZ selects two breaks. For Netherlands, Norway & Turkey, the sequential procedure selects a single break. For U.K. & U.S.A., the sequential procedure selects zero & one break respectively. For all the remaining countries, at least one of the procedures provide evidence against the stability of the long run relationship.

Since the above stability tests also reject the null of coefficient stability when the regression is a spurious one, we still need to confirm the presence of cointegration among the variables. Since we use the cointegration tests as confirmatory tests, for any given country we only consider the tests that correspond to the number of breaks selected by the sequential procedure & information criteria. For example, in the case of Australia, the sequential procedure selects no break while the LWZ criterion selects three breaks. Hence, when verifying cointegration, we only use the usual no-break ADF & PP cointegration tests & the three breaks Arai-Kurozumi test. Table 4 show results for the usual PP & ADF tests of the null hypothesis of no cointegration between saving & investment rates. Again, as in the case of unit root tests, no time trend is included in the cointegrating regression. For Australia & Belgium the null is rejected by the PP tests at the 1% level. The ADF test is not significant for most of the countries. Table 5 presents results for the G-H tests against the alternative of a single break in both intercept & slope (Z_t^*, Z_α^*, ADF_t^*). The results show important differences across countries. The tests for Canada do not seem to support cointegration between the variables (except ADF_t^* which is significant at the 5% level but not at the 1% level). For Finland, Netherlands, Norway, Turkey, Switzerland & Korea, at least one of the G-H tests suggest cointegration at the 1% level. On the other hand, the results for Canada & U.S.A. do not seem to support cointegration among the variables.

As discussed earlier, in order to avoid the problems with the G-H class of tests, we also consider testing the null of cointegration with structural breaks. Table 5 presents results of the single break test together with the estimate of the break date obtained by minimizing the sum of squared residuals. Again, the level of trimming used is 15%. We find that for all countries except Norway, the null is not rejected at the 5% level. This suggests there is some kind of cointegration between the variables (since the test may fail to reject even with a stable cointegrating vector). Finally, Table 6 presents results of cointegration tests allowing for multiple breaks. Results are also reported for other countries for which multiple breaks are selected by at least one of the three procedures. The critical values for the test are then simulated for the corresponding break fractions.[7] The results show that except for Australia, Portugal (at the 5% level only), Italy (at the 10% level only) & Finland, the tests do not reject the null of cointegration.

In order to compare coefficient estimates obtained from a breaks model with those reported in Table 1, Tables 7.1 and 7.2 report estimated parameters & their standard errors under the alternative hypothesis of instability where the number of breaks is selected by the sequential procedure & the information criteria.[8] The coefficient estimates for Australia in a three breaks model show a tendency to increase over time. For Austria, the slope coefficient increases from 0.41 to 0.73 & then falls back to 0.42. Thus the coefficient in each regime is smaller than that obtained from a cointegrating regression without allowing for breaks (0.79 from Table 1). This suggests that ignoring shifts in the long-run relation may overstate the extent of correlation between saving & investment. For Canada, the coefficient estimates in the two regimes are much smaller than the full sample value. In the case of Finland, the estimates of the single break model show that the slope estimate is insignificant in the second regime. As in the case of Austria, the extent of correlation is overstated when not taking into account the shift in the long run relationship (1.04 in Table 1 versus 0.85 & 0.32 in Table 7.1). The pattern of coefficient values for the two breaks model is similar to the case of Austria. For Germany, we estimate both a two breaks model as suggested by LWZ as well as a one break model with the break exogenously imposed at the date of reunification (1990:Q3).[9] Again, the coefficient values in the second & third regimes are much lower than the full sample estimate of 0.55. The result for the exogenous break model is especially interesting in that the coefficient estimate drops from 0.74 in the first regime to 0.05 in the second regime. The results for most of the remaining countries are similar in that the coefficient estimates in more recent regimes are substantially smaller than those in earlier regimes.

[7] The critical values are available upon request.

[8] In Tables 7.1 and 7.2, we evaluate significance using a two-sided t test.

[9] We impose the reunification date as the break date since it provides us with a natural framework to analyze how the saving-investment nexus varies with the size of the country.

Table 7.1: Estimated Regressions under Breaks

	Aus (*L*)	Austria (*S,B,L*)	Can (*L*)	Net (*S*)	Tur (*S,B,L*)	Den (*L*)	Fra (*L*)	Ita (*B,L*)	Jap (*L*)	Spa (*L*)	Swe (*S,B,L*)	Swi (*S,L*)	Bel (*S,B*)
c_1	.21**	.15**	.17**	.01**	.22**	-.11**	.17**	.08**	.01*	.11**	.12*	-.04**	.01**
	(.005)	(.004)	(.007)	(.008)	(.009)	(.008)	(.005)	(.011)	(.004)	(.006)	(.005)	(.006)	(.005)
c_2	.16**	.03**	.12**	.15**	.14**	-.13**	.18**	.21**	.04**	.31**	.02**	-.15**	.03**
	(.007)	(.004)	(.007)	(.004)	(.015)	(.008)	(.005)	(.009)	(.004)	(.006)	(.006)	(.007)	(.003)
c_3	.10**	.13**	-	-	-	.04**	-.01*	.10**	.02**	.40**	.09**	-	-.14**
	(.007)	(.004)				(.004)	(.004)	(.008)	(.003)	(.009)	(.004)		(.004)
c_4	.00	-	-	-	-	-	.19**	.10**	-	.19**	-	-	-
	(.005)						(.006)	(.011)		(.007)			
δ_1	.29	.41**	.23	.98	.13	1.59**	.25	.71	.86**	.57	.13	1.09**	.99*
	(.506)	(.154)	(.382)	(.730)	(.097)	(.461)	(.361)	(1.11)	(.102)	(.445)	(.208)	(.213)	(.427)
δ_2	.40	.73**	.30	.23	.15	1.52**	.15	.16	.84**	-.44	.68**	1.34	.78**
	(.298)	(.159)	(.353)	(.212)	(.178)	(.453)	(.241)	(.545)	(.191)	(.481)	(.248)	(1.26)	(.104)
δ_3	.71	.42#	-	-	-	.61**	1.04**	.51	.85**	-.64	.24	-	1.40**
	(.474)	(.254)				(.292)	(.381)	(.818)	(.091)	(1.46)	(.181)		(.471)
δ_4	1.03**	-	-	-	-	-	-.06	.37	-	.15	-	-	-
	(.290)						(.738)	(1.10)		(.672)			
$\hat{T}_1$	'72:Q3	'75:Q4	'82:Q1	'81:Q1	'00:Q4	'82:Q4	'77:Q3	'75:Q4	'72:Q4	'78:Q3	'88:Q1	'92:Q3	'83:Q2
$\hat{T}_2$	'80:Q2	'87:Q4	-	-	-	'87:Q4	'83:Q1	'82:Q4	'82:Q4	'87:Q2	'92:Q4	-	'93:Q2
$\hat{T}_3$	'87:Q4	-	-	-	-	-	'94:Q2	'92:Q4	-	'91:Q2	-	-	-

Table 7.2: Estimated Regressions under Breaks (Contd.)

	Fin		Ger		Nor		U.S.A.		Por		Kor	
	(S,L)	(B)	(L)	*Exog.*	(S)	(L)	(S)	(B,L)	(B)	(L)	(S)	(L)
c_1	.04**	.19**	.09	.03**	.28**	.28**	.09	.10**	.37**	.38**	.20**	.20**
	(.005)	(.008)	(.006)	(.005)	(.016)	(.019)	(.002)	(.003)	(.011)	(.016)	(.007)	(.007)
c_2	.10**	.06**	.21	.21**	.28**	.60**	.12	.12**	.25**	.27**	-.07**	.17**
	(.010)	(.006)	(.007)	(.009)	(.021)	(.025)	(.002)	(.004)	(.014)	(.010)	(.010)	(.010)
c_3	-	.10**	.14	-	-	.35**	-	.00	.22**	-	-	.003
		(.010)	(.008)			(.022)		(.005)	(.014)			(.011)
c_4	-	-	-	-	-	.28**	-	.14**	.17**	-	-	-.003
						(.017)		(.006)	(.011)			(.010)
δ_1	.85**	.35	.58	.74**	.07*	.09**	.49*	.48	-.54	-.57	.34**	.43*
	(.160)	(.367)	(.564)	(.163)	(.035)	(.028)	(.193)	(.318)	(.443)	(.640)	(.030)	(.026)
δ_2	.32	.74**	.00	.05	-.27	-.92	.42**	.43	-.16	-.14	1.14**	.38**
	(.460)	(.165)	(.302)	(.715)	(.575)	(.700)	(.082)	(.269)	(.185)	(.220)	(.282)	(.146)
δ_3	-	.29	.35	-	-	-.26**	-	1.05	.30	-	-	1.01**
		(.419)	(1.18)			(.676)		(.662)	(.908)			(.340)
δ_4	-	-	-	-	-	-.26	-	.35	.54	-	-	.89**
						(.453)		(.367)	(.982)			(.320)
$\hat{T}_1$	’92:Q4	’77:Q4	’73:Q1	’90:Q3	’89:Q4	’73:Q4	’76:Q4	’77:Q1	’83:Q4	’83:Q3	’89:Q4	’81:Q2
$\hat{T}_2$	-	’92:Q4	’90:Q4	-	-	’80:Q3	-	’88:Q4	’87:Q4	-	-	’90:Q1
$\hat{T}_3$	-	-	-	-	-	’89:Q4	-	’98:Q3	’91:Q3	-	-	’97:Q3

Overall, for most countries saving-investment correlations are higher during the '70s suggestive of disintegration tendencies in the wake of the collapse of the Bretton Woods system & the ensuing confusion in international financial markets. National policy responses to the first oil shock may also have played a part in generating an autarkic tendency in capital markets. Both events could be associated with a structural shift in the system due to, say, perceived increases in the risks associated with international lending. A general decline in coefficients is evident through the early/mid '80s heralding the return of integrative forces & increased openness of economies.

5 Discussion

It is important to emphasize that the finding of cointegration *per se* does not have any implications for the extent of capital mobility. The standard intertemporal open economy macroeconomic theory implies that saving & investment should be cointegrated even if capital is perfectly mobile. A country needs to rely on its current account balance in order to repay its external debt. Therefore, unless a country's intertemporal budget constraint is violated, no investment-saving gap can remain permanent. Hence, testing for unit cointegration between saving & investment essentially amounts to testing the solvency of the economy. (see Levy, 2000 & Coakley & Kulasi, 1997). On the contrary, this paper is primarily aimed at investigating to what extent this long run correlation varies with the degree of observed international capital mobility.

One reason why saving & investment shares might be correlated even in the presence of capital mobility is the effect of country size. The country size argument can be found in two versions. Harberger (1980) argues that as countries become larger, they become more diversified & the need to borrow from abroad in the event of a shock declines. Since the original results of Feldstein & Horioka are based on a sample that includes some very large countries, the Harberger argument could be an explanation for the high saving-investment correlations. The second version of the country size argument related the size of a country to its influence on the world interest rate. If a country is big enough to affect the world interest rate an increase in national saving would lower the world interest rate & increase investment in that country. Saving & investment would be correlated although perfect capital mobility prevails.

In their original work, Feldstein & Horioka (1980) examined the possibility that the link between domestic investment & domestic saving varies with the size of the economy. They used the logarithm of GDP to measure size so that the variance of the variable would not be dominated by the few largest observations. However, they did not find any evidence that the saving-investment correlation varied with the size of the economy. Feldstein & Horioka (1980) also investigated the possibility that the saving-investment correlation varies with the degree of openness of the economy. The measure of openness used is the share of trade in GDP as measured by the sum of exports & imports per dollar of GDP. They found that the saving-investment link does not seem to bear any relation to the importance of international trade.

Baxter & Crucini (1993) present rankings of 8 OECD countries in terms of their saving-investment correlation as well as their size as measured by GNP. On the basis of these rankings which they also justify using a theoretical model, they argue that larger countries have larger correlations while correlations are still substantial for smaller countries. Our results stand in stark contrast to those in Baxter & Crucini (1993). In particular, the presence of breaks & the relatively smaller coefficient values in the last regime indicate that the saving-investment association does not bear any relationship with the size of the economy. For instance, when the number of breaks is selected using the sequential procedure, the R^2 values do not exceed 50% for any of the countries which experience breaks. As another example, consider the case of Germany. When we estimate the model imposing the exogenous break at the time of reunification, the coefficient estimate in the post-reunification regime is substantially smaller than that in the pre-reunification regime. (0.74 versus 0.05). If the saving-investment association were to vary positively with the size of the country, we would at least expect an increase in the value of the slope estimate from the first to the second regime.

To further investigate whether the saving-investment association bears a relationship to country size, we use the 2005 World Bank estimates of total GDP as well as PPP GDP.[10] Table 8 presents a comparison of rankings of the countries with respect to the R^2 in the last regime & these two measures of country size.[11] We consider the sequential & LWZ procedures to select the number of breaks. If a procedure does not find any evidence of instability for any particular country, we use the full-sample R^2 value for that country. U.K. is excluded since we do not find any evidence of cointegration for this coun-

[10] The data is available as "Quick Reference Tables" on the World Bank website. Purchasing power parity (PPP) conversion factors take into account differences in the relative prices of goods and services—particularly non-tradables—and therefore provide a better overall measure of the real value of output produced by an economy compared to other economies.

[11] The rankings for size are from largest to smallest and for R^2 is that from highest to lowest.

Table 8: Rankings of Countries by Size and R^2

Ctry	(S)	Tot	PPP	Ctry	(S)	Tot	PPP	Ctry	(L)	Tot	PPP	Ctry	(L)	Tot	PPP
Aus	2	10	10	Fra	3	4	4	Aus	2	10	10	Fra	3	4	4
Austria	15	16	15	Mex	20	9	7	Austria	15	15	14	Mex	19	9	7
Can	12	7	8	Ita	4	5	5	Can	12	7	8	Ita	4	5	5
Fin	11	19	20	Por	15	20	17	Fin	11	18	19	Por	15	19	16
Ger	5	3	3	Jap	1	2	2	Ger	5	3	3	Jap	1	2	2
Net	17	11	12	Spa	9	6	6	Net	-	-	-	Spa	9	6	6
Nor	12	17	18	Swe	14	15	14	Nor	12	16	17	Swe	14	14	13
Tur	18	14	11	Swi	7	13	16	Tur	17	13	11	Swi	7	12	15
U.S.A.	7	1	1	Bel	9	12	13	U.S.A.	7	1	1	Bel	9	11	12
Den	19	18	19	Kor	6	8	9	Den	18	17	18	Kor	6	8	9

Table 9: Trade Shares and R^2

(S)					(B)					(L)				
Ctry	A		B		Ctry	A		B		Ctry	A		B	
	R^2	T.Sh	R^2	T.Sh		R^2	T.Sh	R^2	T.Sh		R^2	T.Sh	R^2	T.Sh
Austria	.32	.35	.16	.37	Austria	.32	.35	.16	.37	Aus	.34	.32	.73	.38
Fin	.56	.54	.22	.66	Fin	.67	.55	.22	.66	Austria	.32	.35	.16	.37
Net	.68	.97	.11	1.08	Tur	.14	.42	.10	.63	Can	.06	.21	.13	.33
Nor	.02	.80	.21	.72	U.S.A.	.76	.22	.50	.25	Fin	.56	.54	.22	.66
Tur	.14	.42	.10	.63	Ita	.32	.40	.37	.45	Ger	.00	.56	.22	.50
U.S.A.	.54	.11	.41	.21	Por	.30	.67	.21	.62	Nor	.14	.74	.21	.72
Swe	.55	.58	.20	.78	Spa	.24	.37	.00	.44	Tur	.14	.42	.10	.63
Swi	.85	.68	.41	.77	Swe	.55	.58	.20	.78	U.S.A.	.76	.22	.50	.25
Bel	.60	1.36	.32	1.36	Bel	.40	1.36	.32	1.36	Den	.79	.68	.13	.76
Kor	.09	.52	.44	.68						Fra	.77	.43	.32	.47
										Ita	.32	.40	.37	.45
										Por	.62	.67	.62	.64
										Jap	.76	.13	.92	.11
										Spa	.24	.37	.00	.44
										Swe	.55	.58	.20	.78
										Swi	.85	.68	.41	.77
										Kor	.49	.57	.60	.77

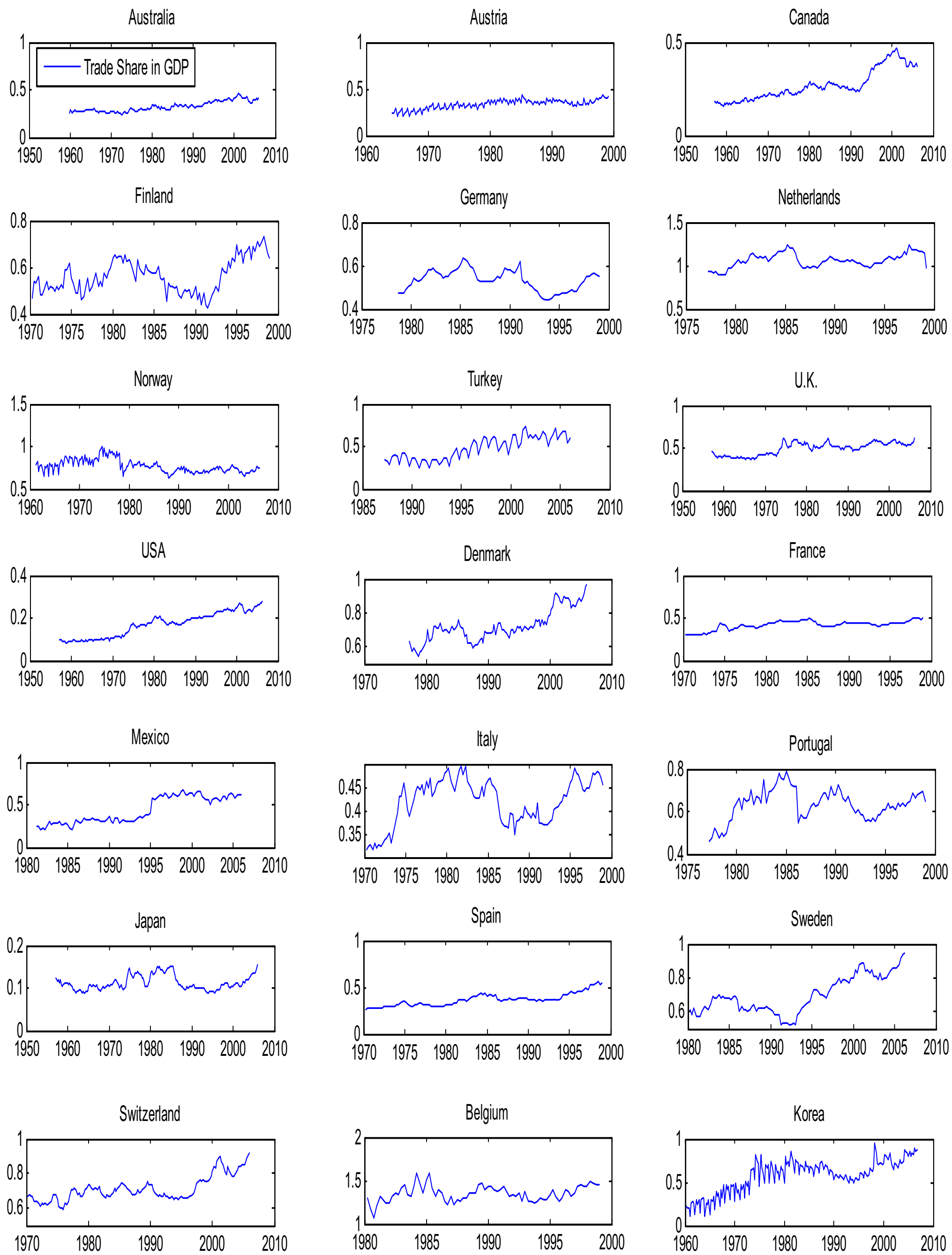

Figure 3: Share of Trade in GDP

try. Also, the information criterion LWZ selects more than three breaks for Netherlands & U.K.. We thus exclude these two countries when computing the rankings based on this criterion. The results do not suggest any clear correspondence between the saving-investment link & country size. For example, Australia is ranked in the second place using both the sequential procedure & the information criterion LWZ while its ranking in terms of country size is much lower. As another example, U.S.A. is ranked in the seventh place by both the sequential procedure & the LWZ criterion respectively while it ranks first in terms of country size. Thus, our results do not support the claim that saving investment correlations are larger in larger countries.

Contrary to Feldstein & Horioka (1980), however, we find that the correlation between these two macroeconomic variables vary closely with the degree of openness of the economy. To illustrate our point, we use the same measure of openness, namely, the share of trade in GDP. Figure 3 plot these shares for the 21 OECD countries. It is evident from the figure that the share of trade is generally higher in the later periods. This is consistent with integration of world capital markets & a high degree of international diversification in the 1990s. To measure precisely the association with the degree of openness, we compute the R^2 in the last two regimes for each country & each procedure for which we select at least one break. We then compare these values with the average trade shares in those regimes. The results are presented in Table 9. The final regime is labelled regime B while the preceding regime is labelled regime A. We see that for all countries except Germany, Norway, Portugal & Japan, the trade share in the last regime is higher than that in the preceding regime. With the exception of Australia & Canada, we see that considering the last two regimes, the R^2 bears a negative relationship with the degree of openness in that a higher trade share corresponds to a lower R^2 & vice-versa. The average R^2 (across countries) in regime A when the number of breaks is selected using the sequential procedure is 0.44 while the corresponding average for regime B is only 0.26. When the number of breaks is selected using the information criterion LWZ, the average R^2 falls from 0.45 to 0.34 while when it is selected using the information criterion BIC, the drop is sharpest from 0.41 to 0.23. We believe this provides strong evidence in favour of a close association between the saving-investment correlation & the degree of openness. Note that U.K. & Mexico are excluded from this table since the sequential procedure selects no break & the information criteria select too many breaks. However, as is evident from Figure 3, the trade share for U.K. has remained quite stable over the entire sample period. The association between the saving-investment link & openness could thus be a potential explanation of the fact that the sequential procedure does not provide any evidence of instability in the long run relationship between the variables for this particular country. For Austria & Finland, the estimated coefficients from a two breaks model are such that the value of the slope jumps to a higher value during the '70s & subsequently returns to a lower value (comparable to its value in the first regime) in the late '80s/early '90s. Given that the decade of the '70s was a turbulent period for the world economy as a whole (& especially for small countries like Austria & Finland which are likely to be more susceptible to such turbulences), the extent of international diversification was relatively low which can explain the higher coefficient value in the second regime.

6 Conclusion

This paper analyzes the well known Feldstein-Horioka saving-investment puzzle from a time series perspective as opposed to a cross section/panel data perspective. In particular, it is shown, using a sample of 21 OECD countries, that saving & investment rates are nonstationary over the sample period considered. Further, the nature of the long run relationship is analyzed using a battery of cointegration tests which include existing tests as well as their extended versions to allow for multiple breaks. These tests are used in conjunction with tests of stability recently proposed in Kejriwal & Perron (2006a). It is found that for all countries except the U.K. & Mexico, there is some evidence of instability in the cointegrating relationship. Our estimates of the cointegrating vector over different regimes suggest a strong linkage between the saving-investment correlation & the degree of openness, as measured by the share of trade in GDP. Hence, our empirical results are consistent with the recent evidence on international diversification & do not seem to suggest the existence of a puzzle as advocated by Feldstein & Horioka (1980). We thus argue that while Feldstein & Horioka's basic idea that the saving-investment correlation contains information about international capital mobility is correct, our time series analysis shows that the extent of such correlation may be overstated due to incorrect specification of the regression model. Finally, our results also clearly refute the commonly held view that the saving-investment association bears a close relationship with the size of the country.

Technical Appendix

(A) Description of Unit Root Tests: For any series $\{x_t\}_{t=0}^{T}$, define $(x_0^{\bar{\alpha}}, x_t^{\bar{\alpha}}) = (x_0,\ (1-\bar{\alpha}L)x_t),\ t = 1, ..., T,\ x^{\bar{\alpha}} = (x_1^{\bar{\alpha}}, ..., x_T^{\bar{\alpha}})'$ for some chosen $\bar{\alpha} = 1 + \bar{c}/T$. Let $z_t = \{1\}$. The GLS detrended series is defined as

$$\tilde{y}_t = y_t - \hat{\psi}' z_t$$

where

$$\hat{\psi} = \arg\min_{\psi} \, (y^{\bar{\alpha}} - \psi' z^{\bar{\alpha}})'(y^{\bar{\alpha}} - \psi' z^{\bar{\alpha}})$$

Ng & Perron (2001) propose the following tests based on GLS detrended data :

$$\begin{aligned}
MZ_{\alpha}^{GLS} &= (T^{-1}\tilde{y}_T^2 - s_{AR}^2)(2T^{-2}\sum_{t=1}^{T}\tilde{y}_{t-1}^2)^{-1} \\
MSB^{GLS} &= (T^{-2}\sum_{t=1}^{T}\tilde{y}_{t-1}^2/s_{AR}^2)^{1/2} \\
MZ_{t}^{GLS} &= MZ_{\alpha}^{GLS} \times MSB^{GLS} \\
MP_{T}^{GLS} &= (\bar{c}^2 T^{-2}\sum_{t=1}^{T}\tilde{y}_{t-1}^2 - \bar{c}T^{-1}\tilde{y}_T^2)/s_{AR}^2
\end{aligned}$$

where

$$s_{AR}^2 = \hat{\sigma}_k^2/(1 - \hat{\beta}(1))^2 \tag{6}$$

In (3), $\hat{\beta}(1) = \sum_{i=1}^{k}\hat{\beta}_i$, $\hat{\sigma}_k^2 = (T-k)^{-1}\sum_{t=k+1}^{T}\hat{e}_{tk}^2$, with $\hat{\beta}_i$ & $\{\hat{e}_{tk}\}$ obtained from the following regression estimated by OLS:

$$\Delta\tilde{y}_t = \psi + \beta_0\tilde{y}_{t-1} + \sum_{j=1}^{k}\beta_j\Delta\tilde{y}_{t-j} + e_{tk}$$

Ng & Perron (2001) also propose a class of Modified Information Criteria (MIC) that selects k as $k_{mic} = \arg\min_k MIC(k)$ where

$$MIC(k) = \ln(\hat{\sigma}_k^2) + \frac{C_T(\tau_T(k) + k)}{T - k_{\max}}, \; C_T > 0 \text{ \& } C_T/T \to 0 \text{ as } T \to \infty$$

where $\tau_T(k) = (\hat{\sigma}_k^2)^{-1}\hat{\beta}_0^2 \sum_{t=k_{\max}+1}^{T}\tilde{y}_{t-1}^2$ & $\hat{\sigma}_k^2 = (T-k_{\max})^{-1}\sum_{t=k_{\max}+1}^{T}\hat{e}_{tk}^2$. The $MAIC$ obtains with $C_T = 2$ while the $MBIC$ obtains with $C_T = \ln(T-k_{\max})$.

(B) Proof of Proposition: We can write (4) in matrix form as

$$Y = \bar{W}\gamma + U$$

where $W = (w_1, ..., w_T)'$, $w_t = (1, z'_{2t})'$, $\gamma = (c_1, \delta'_1, ..., c_{m+1}, \delta'_{m+1})'$, & $\bar{W}$ is the matrix which diagonally partitions W at the $k-$partition $(T_1, ..., T_k)$, that is, $\bar{W}$ =diag$(W_1, ..., W_{k+1})$ with $W_i = (w_{T_{i-1}+1}, ..., w_{T_i})'$. We write $\bar{W} = (\bar{w}_1, ..., \bar{w}_T)'$. Let the residuals be denoted $\hat{u}_t$ $(t = 1, .., T)$. Define the matrices

$$D_{1T} = diag(T^{-1},\ T^{-3/2}I_q), D_{2T} = diag(T^{-1/2},\ T^{-1}I_q)$$

Then we have

$$T^{-1/2}\sum_{j=1}^{t}\hat{u}_j = T^{-1/2}\sum_{j=1}^{t}(y_j - \bar{w}'_j\hat{\gamma})$$

$$= T^{-1/2}\sum_{j=1}^{t}(u_j - \bar{w}'_j(\hat{\gamma} - \gamma))$$

$$\begin{aligned} &= T^{-1/2}\sum_{j=1}^{t}(u_j - \bar{w}'_j(\sum_{f=1}^{T}\bar{w}_f\bar{w}'_f)^{-1}\sum_{f=1}^{T}\bar{w}_f u_f) \\ &= T^{-1/2}\sum_{j=1}^{t}u_j - g(t) \end{aligned}$$

where

$$g(t) = \sum_{j=1}^{t}w'_j D_{1T}\{\sum_{i=1}^{k+1}(D_{2T}(\sum_{f=1}^{T}w_f w'_f)D_{2T})^{-1}(D_{2T}\sum_{f=1}^{T}w_f u_f)I(T_{i-1} < j \leq T_i)\}.$$

Hence, we have

$$T^{-1/2}\sum_{j=1}^{t}\hat{u}_j \Rightarrow W_1(r) - G(r)$$

where $G(r)$ is as defined in the text with $W(s) = (1, W_2(r)')'$, W_2 being a vector of q independent Wiener processes & W_1 is a one dimensional Wiener process independent of W_2. The result follows directly by application of the continuous mapping theorem.

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