Chapter 4

IS THE FELDSTEIN-HORIOKA PUZZLE REALLY A PUZZLE?

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Abstract: Using the framework of a dynamic intertemporal optimization model of an open economy, it is shown that the long-run investment-saving correlation follows directly from the economy's dynamic budget constraint and this does not depend on the degree of international capital mobility. Therefore, unless the budget constraint is violated, the time series of investment and saving should be cointegrated, and this should be true for any degree of capital mobility. Using an improved econometric technique, which encompasses the tests used by previous authors and avoids some of the pitfalls associated with their tests, I show that their conflicting findings can be explained by a simple but important, omitted variables problem. Using annual and quarterly post-war U.S. data, I find that investment and saving are cointegrated in levels as well as in rates, regardless of the time period considered, as predicted by the model.

Keywords: Capital Mobility, Investment-Saving Correlation, Dynamic Budget Constraint, Integration and Cointegration, Omitted Variables

1. INTRODUCTION

There is strong empirical evidence that domestic investment (I) and national saving (S) are correlated.[1] Much of the evidence is based on cross-section regressions of multi-year average data and therefore, this is considered to be a long-run phenomenon. This finding, also known as the

Feldstein-Horioka (1980) puzzle, has received significant attention because Feldstein and Horioka interpret it as evidence of low international capital mobility. In a closed economy, investment must be financed by saving. In an open economy, however, some of the investment may be financed by foreign saving and therefore, saving and investment could move independently from each other. Thus, the high *I-S* correlation, Feldstein and Horioka argue, suggests that capital might not be mobile. This conclusion, however, is in contrast with the deregulation of capital markets and increased integration of world financial markets in the last 30 years. Also, studies measuring capital mobility directly using PPP and various interest parity conditions, conclude that capital is very mobile.[2]

Knowing the true degree of capital mobility is important for several reasons. For example, the effect of fiscal policy crucially depends on the extent of capital mobility. In addition, an economy's access to capital markets can reduce the cost of adjustment to external shocks. Also, capital mobility determines the rate at which incomes converge. Further, perfect capital mobility is often assumed to hold in macroeconomic models. Capital immobility would call into question this common practice.[3]

The existing time series studies of *I-S* comovement report conflicting results. For example, Miller (1988) finds that saving and investment in the US are cointegrated during the fixed, but not during the flexible exchange-rate period, and concludes that increased capital mobility since the 1970s may have severed the *I-S* link. Gulley (1992) uses an improved test and finds that saving and investment are not cointegrated in either period. Otto and Wirjanto (1989) also conclude that saving and investment in the U.S. are not cointegrated. Alyousha and Tsoukis (in this volume) use data that cover a longer time horizon but they also find no cointegration.

This chapter claims that there is nothing puzzling in the Feldstein and Horioka's finding. The neoclassical growth model predicts that, in the steady state, investment and saving would be proportional to output.[4] It would be puzzling, therefore, if we did *not* find high *I-S* correlation.[5]

Most optimization-based dynamic models of open economy also predict that investment and saving should be correlated in the long run. Optimizing individuals face intertemporal budget constraint, which implies that, in the long run, current account balances should add up to zero as current account surpluses or deficits cannot be sustained forever. Thus, in the long run, investment and saving would be correlated, regardless of the degree of capital mobility, as long as the intertemporal budget constraint is not violated. A test of *I-S* cointegration, therefore, is merely a test of country's economic solvency.

It follows that the time series of investment and saving should be cointegrated, and this would be true for any degree of capital mobility. Using an improved econometric technique which encompasses the tests used by the above authors and avoids some of the pitfalls associated with their tests, I show that their conflicting findings can be explained by a simple, but important, omitted variables problem. In particular, using annual and quarterly post-war U.S. data, I demonstrate that even if investment and saving are not cointegrated in a bivariate setup, they are cointegrated when output is added to the system. In order to allow for the possibility of structural breaks in the *I-S* relationship, I consider the entire post-war sample period as well as its several sub-periods. It turns out that the cointegration finding is robust regardless of the time period considered.

Thus, the U.S. data do not violate the intertemporal budget constraint and so the U.S. economy is solvent. The main conclusion is that the observed long-run *I-S* correlation cannot be useful in measuring the degree of long-term capital mobility.

The chapter is organized as follows. In the next section I derive the long-run implication of the intertemporal budget constraint of an open economy and discuss its interpretation in the context of the empirical findings reported below. In section 3, I discuss omitted variables problem in cointegration tests. In section 4, I discuss the integration tests and present their results. The cointegration test results are reported in section 5. The paper ends with a brief summary and concluding remarks in section 6.

2. INTERTEMPORAL BUDGET CONSTRAINT

Consider a dynamic optimization model of an open economy with a budget constraint of the form:

$$\frac{dB_t}{dt} = \rho_t B_t + C_t + G_t + I_t - Y_t, \qquad (1)$$

where ρ is time varying world interest rate, B is foreign debt, C is consumption, G is government expenditure, I is investment, and Y is output. According to (1), the change in foreign debt equals spending minus production, where spending includes interest payments on the existing debt. The idea behind this constraint is that an economy may borrow from abroad to pay for excess spending, or it may lend to a foreign country to accommodate excess production. Thus, world capital markets enable the economy to accommodate temporary imbalances between production and spending.

It is well known that the intertemporal budget constraint given in (1) is actually a nonhomogenous differential equation. Integrating forward yields:

$$B_t = A\psi_t^{-1} + \psi_t^{-1}\int_t^{\infty} \psi_s (Y_s - C_s - G_s - I_s)\, ds, \qquad (2)$$

where A is set to zero, and $\ln\psi_t = -\int_0^t \rho_s ds$, where ψ_t is the discount factor applied to the returns of the time t-period into the future. In a similar fashion, $\ln\psi_s = -\int_0^s \rho_v dv = -\int_t^{t+s} \rho_v dv$, which is used in deriving (2). The discount factor $\psi_t^{-1}\psi_s$ gives the time t-value of a dollar to be delivered at time s.

Now let us assume that the $\lim_{t\to\infty}(\psi_t B_t) = 0$, which is the non-Ponzi game condition. This prevents the representative agent from incurring ever-increasing debt by continuously borrowing without a limit. At the same time, however, the assumption does not impede the agent's ability to incur a temporary debt to accommodate temporary imbalance between production and spending.

The above budget constraint can be used to relate the long-run *I-S* comovement to current account stationarity. Assume $\rho_s = \rho, \forall s$. Then ψ_t becomes the standard continuous-time discount factor with constant interest rate, $\ln\psi_t = -\int_0^t \rho ds = -\rho t$. In this case (2) can be rewritten as:

$$e^{-\rho t} B_t = \int_t^{\infty} e^{-\rho s}(Y_s - C_s - G_s - I_s)\, ds, \qquad (3)$$

which, using the fact that $e^{-\rho t} = \int_t^{\infty} \rho e^{-\rho s} ds$ can be further rewritten as

$$\int_t^{\infty} e^{-\rho s}(Y_s - \rho B_s - C_s - G_s - I_s)\, ds = 0, \qquad (4)$$

where $Y_s - \rho B_s$ denotes the net income of domestic residents, the *GNP*. But $S = GNP - C - G = Y - \rho B - C - G$, which follows from the national income accounting. Therefore, in (4), the term in parentheses equals $S - I$, which in turn equals the current-account deficit.

Thus, a long-run *I-S* correlation is equivalent to a stationarity of current account deficit. Therefore, if investment and saving are cointegrated, it is an indicator of the country's economic solvency. As Obstfeld (1991), Alyousha and Tsoukis (in this volume) and Coakley, et al. (1996) emphasize, in a model with a variable real interest rate, stationarity of current account is sufficient for external solvency. The implication of (4), however, is that in a model with a constant real interest rate, stationarity of current account is both necessary and sufficient for economic solvency.

3. OMITTED VARIABLES IN COINTEGRATION

Since investment and saving tend to be non-stationary, Miller (1988), Otto and Wirjanto (1989), and Gulley (1992) use the cointegration methodology to study the *I-S* relationship in the post-war U.S. All three use Engle and Granger's (1987) two-step estimation method, but report conflicting findings. Miller (1988) finds that the series are cointegrated prior to 1971, during the fixed exchange rate period, but not after 1971, during the flexible exchange rate regime. Otto and Wirjanto (1989) and Gulley (1992), however, find that the series are not cointegrated in either period.

Because investment and saving must be cointegrated, these conflicting findings may be due to an omitted variable. Consider a situation where y, x_1, and x_2 are all $I(1)$, but their linear combination is $I(0)$. In other words, I assume that the time series of y, x_1, and x_2 are cointegrated, which means that $y = \beta_1 x_1 + \beta_2 x_2 + \varepsilon$, where $\varepsilon \sim I(0)$. Now, suppose that that we inadvertently omit x_2 and run $y = \beta_1 x_1 + \mu$. Since $\mu = (\beta_2 x_2 + \varepsilon) \sim I(1)$, we would mistakenly conclude that y and x_1 are not cointegrated.

This example suggests the possibility that the conflicting results reported in the above studies may be caused by omission of some important variable. According to the neoclassical growth model, a natural candidate for a missing variable is output because in that model, investment and saving are proportional to output.

4. INTEGRATION TEST RESULTS

To test for stationarity, Miller (1988) uses the Augmented Dickey-Fuller (ADF) unit root test:

$$\Delta x_t = \gamma x_{t-1} + \sum_{i=1}^{4} \phi_i \Delta x_{t-i} + \varepsilon_t \tag{5}$$

However, Gulley (1992) correctly claims that the exclusion of the constant is appropriate only if the mean of the series is zero, which is not the case for saving or investment. Therefore he modifies (5) by adding a constant,

$$\Delta x_t = \alpha_1 + \gamma x_{t-1} + \sum_{i=1}^{4} \phi_i \Delta x_{t-i} + \varepsilon_t \qquad (6)$$

and tests for $\gamma = 0$.

However, this version of the ADF test is not problem-free either. The reason is that the tabulated distribution of the unit root test statistic for version (6) depends crucially on the assumption that $\alpha_1 = 0$. That is, it has the Dickey-Fuller distribution only when there is no drift term in the data-generating process of x_t. If the true $\alpha_1 \neq 0$, then the statistic for testing the null hypothesis $\gamma = 0$ is asymptotically distributed as $N(0,1)$, and, in finite samples, its distribution may or may not be well approximated by the Dickey-Fuller distribution.[6] Therefore, if the drift parameters in the data-generating processes of investment and saving are non-zero, then using version (6) of the test is inappropriate.

To avoid the dependence of the distribution of the test statistic on the value of α_1, MacKinnon (1991) suggests adding a linear time trend to (6),

$$\Delta x_t = \alpha_0 t + \alpha_1 + \gamma x_{t-1} + \sum_{i=1}^{4} \phi_i \Delta x_{t-i} + \varepsilon_t, \qquad (7)$$

under the assumption that there is no trend in the data-generating process.

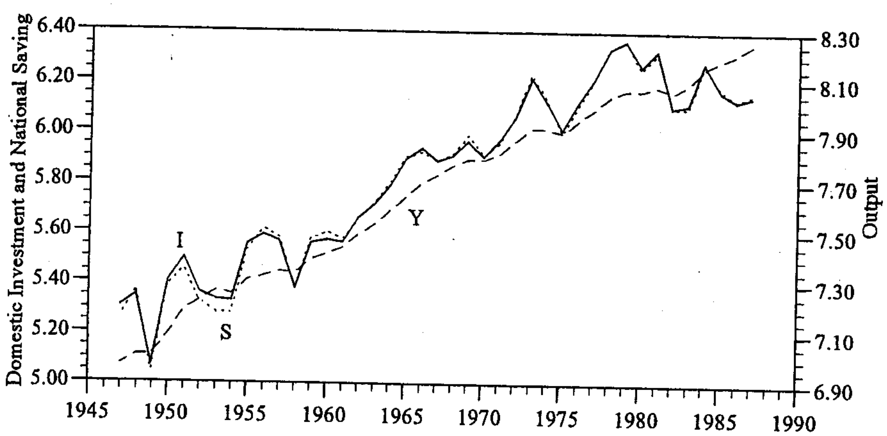

Figure 1. Investment, Saving, and Output in Levels, Annual Data, 1947–87

I use (7) to examine the unit-root properties of the time series of saving, investment, and output. Along with (7), I have used Box-Pierce, Ljung-Box, and Lagrange Multiplier tests (not shown to save space) to verify that the error terms in the unit-root test regressions are not serially correlated.

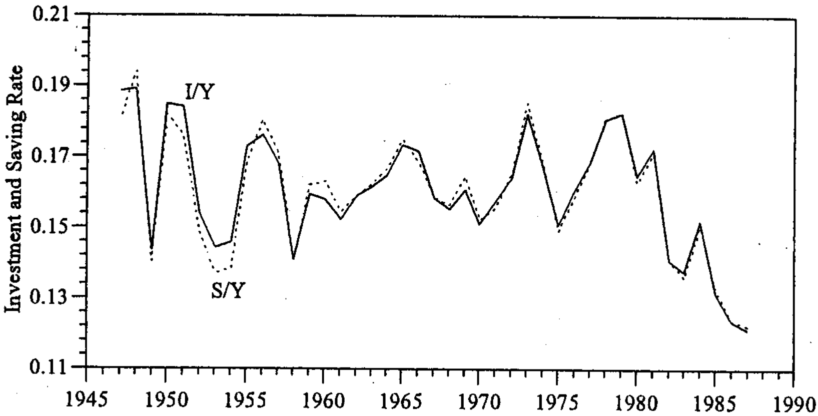

Figure 2. Investment and Saving Rates, Annual Data, 1947–87

I use quarterly and annual data for 1947–1987. The quarterly data are identical to those used by Miller (1988), Otto and Wirjanto (1989), and Gulley (1992). I study the *I-S* relationship both, in levels and in rates.[7]

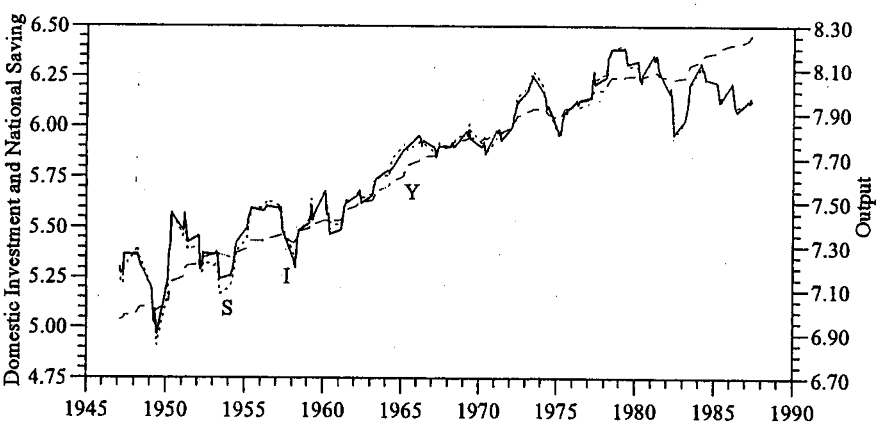

Figure 3. Investment, Saving, and Output in Levels, Quarterly Data, 1947:1–87:3

The source of the data on national saving, domestic investment, and output is the US-NIPA Tables of the Bureau of Economic Analysis.

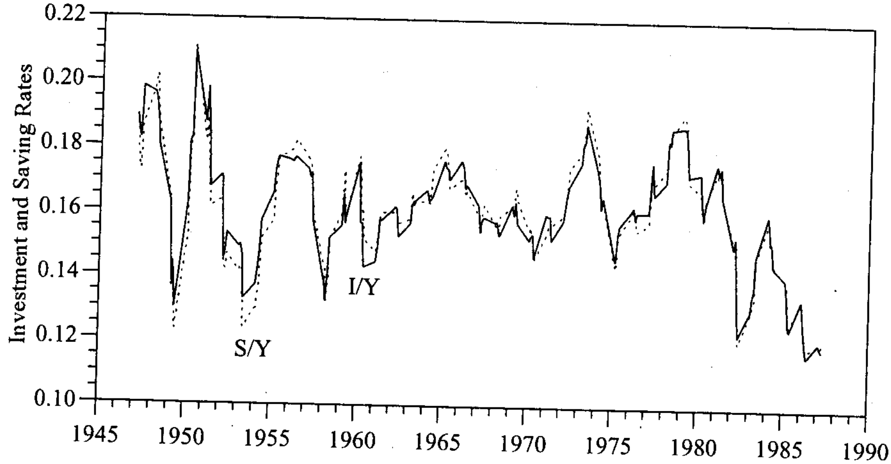

Figure 4. Investment and Saving Rates, Quarterly Data, 1947:1–87:3

In addition to the entire sample period 1947:1–87:3, I also examine its three sub-periods, 1947:1–71:2, 1971:3–87:3, and 1980:1–87:3.

The 1947:1–87:3 sample period was chosen to match the sample periods used by Miller (1988), Otto and Wirjanto (1989), and Gulley (1992).[8] The first two sub-periods, 1947:1–71:2 and 1971:3–87:3, correspond to the fixed and flexible exchange rate regimes, respectively.[9] The last sub-period, 1980:1–87:3, is examined to see whether the large capital inflow into the U.S. during the Reagan administration altered the *I*-*S* relationship.

The annual series measured in levels (log) and as a fraction of output (that is, the investment and saving rates) are displayed in Figures 1 and 2, respectively. Similarly, the quarterly series measured in levels (log) and as a fraction of output are shown in Figures 3 and 4, respectively.

I present the integration test results in Tables 1 and 2. The ADF test statistics indicate that the *S*, *I*, and *Y* series are $I(1)$ when measured in levels. When differenced, all three series appear to be $I(0)$. This is true for both the annual (Table 1) as well as the quarterly data (Table 2). When measured in rates, saving and investment appear to be $I(1)$ during the 1971:3–87:3 and 1980:1–87:3 periods.[10] In what follows, therefore, I treat them as $I(1)$.

Table 1. Unit-Root ADF Test of Investment, Saving, and Output: Annual Data

Period	Series	Level	First Difference (Δ)
1947–87	I	−1.47	−3.57**
(n = 41)	S	−1.59	−3.71
	Y	−1.22	−4.01**
	I/Y	−3.53**	
	S/Y	−3.38**	

Notes: Superscripts *, **, and *** in all tables indicate statistical significance at 1%, 5%, and 10%, respectively. The corresponding MacKinnon (1991) critical values for the ADF test statistics are −4.19, −3.52, and −3.19, respectively. Miller (1988), Otto and Wirjanto (1989), and Gulley (1992) do not use annual data.

Table 2. Unit-Root ADF Tests of Investment, Saving, and Output: Quarterly Data

Period	Series	Miller	Otto and Wirjanto	Gulley	Levy Level	Levy Difference
1947:1–87:3	I		−3.98**	−0.86	−3.33	−6.87*
(n = 163)	S		−3.83**	−1.28	−3.40	−6.90*
	Y				−2.03	−5.51*
	I/Y	−0.14		−4.88**	−3.77**	
	S/Y	−0.77		−3.09**	−3.75**	
1947:1–71:2	I		−3.54**	−0.46	−3.22	−5.48*
(n = 98)	S		−2.54	−0.37	−3.41	−5.19*
	Y				−1.35	−4.48*
	I/Y	−0.04		−3.91**	−4.25*	
	S/Y	−0.40		−3.30**	−4.11*	
1971:3–87:3	I		−3.11	−1.94	−2.66	−4.39*
(n = 65)	S		−3.25	−2.78	−2.76	−4.49*
	Y				−2.43	−3.32***
	I/Y	−0.37		−3.00**	−2.43	−4.43*
	S/Y	−0.74		−1.22	−2.58	−4.44*
1980:1–87:3	I				−2.99	−3.57**
(n = 31)	S				−2.67	−3.63**
	Y				−1.75	−2.73
	I/Y				−3.53***	−3.70**
	S/Y				−3.16	−3.70**

Notes: The corresponding critical values of MacKinnon (1991) for Levy's ADF test statistics are −4.01, −3.43, and −3.14 for 1947:1–87:3, −4.05, −3.45, and −3.15 for 1947:1–71:2, −4.10, −3.47, and −3.16 for 1971:3–87:3, and −4.28, −3.56, and −3.21 for 1980:1–87:3, respectively. Otto and Wirjanto's (1989) sample begins with 1956:1. The ADF statistic values for Miller (1988), Otto and Wirjanto (1989), and Gulley (1992) are taken from the respective studies.

5. COINTEGRATION TEST RESULTS

I use Johansen's (1988) maximum likelihood (ML) method, which is superior to the Engle-Granger (1987) two-step method used by the above authors. In addition to the inferior statistical properties of its estimators, the Engle-Granger method has the disadvantage that for estimating a cointegration relationship, some kind of normalization is necessary. Practical applications have shown that the results can be very sensitive to the normalization chosen. Johansen's method treats all variables as endogenous, thereby avoiding the problem of choosing a normalization altogether.

Johansen (1988) offers two tests for estimating the number of cointegrating vectors. The first is called *maximal eigenvalue test*, and is given by the test statistic $\lambda_{\max} = -n\log(1-\hat{\lambda}_r)$, where n is the number of observations, and $\hat{\lambda}_r$ is the r^{th} eigenvalue to be determined by solving the determinantal equation associated with the residual product moment matrix constructed using the residuals' matrices.

The maximal eigenvalue test is designed to test $H(r-1)$ against $H(r)$. That is, the null hypothesis is that there are $(r-1)$ cointegrating vectors against the alternative r.

The second test, called the *trace test*, is designed to test the null $H(r)$ against the alternative $H(m)$, where $r < m$. The trace test statistic is given by $J_T = -n\sum_{i=r+1}^{m}\log(1-\hat{\lambda}_i)$.

The cointegration test results are presented in Tables 3–7. In estimating the cointegration vectors, I used VAR(4). It is not known a priori whether the true data-generating process contains a deterministic trend or not. I, therefore, conduct the cointegration tests under both options. The test statistics are identical under both assumptions; only the critical values differ.

In Johansen's framework the number of cointegrating vectors is determined sequentially. We start with the hypothesis that there are no cointegrating relations, that is, $r=0$, where r denotes the number of cointegrating relationships. We continue only if this hypothesis is rejected. In this case, we test the hypothesis that there is at most one cointegrating vector, $r \le 1$, and so on. The test results can be interpreted in favor of cointegration only if $0 < r < m$, where m is the number of variables in $\mathbf{x}_t$. Full rank, that is $r = m$, only indicates that the data vector process $\mathbf{x}_t$ is stationary. If $r=0$, then the matrix Π, which is the matrix of the coefficients on the variables $\mathbf{x}_{t-p}$ in the first-differenced VAR model, is the null matrix. In that case, the model becomes a traditional differenced VAR system.

Table 3. Cointegration Test: Annual Data, 1947–87 (n = 41)

Variables	Test	H_0	H_1	Test Statistic	Critical Value			
					Trend in DGP		No Trend in DGP	
					95%	90%	95%	90%
I, S	λ_{max}	$r=0$	$r=1$	6.56	14.06	12.07	14.90	12.91
		$r\leq 1$	$r=2$	1.69	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\geq 1$	8.26	15.41	13.32	17.95	15.66
		$r\leq 1$	$r=2$	1.69	3.76	2.68	8.17	6.50
I, S, Y	λ_{max}	$r=0$	$r=1$	20.18***	20.96	18.59	21.07	18.90
		$r\leq 1$	$r=2$	6.23	14.06	12.07	14.90	12.91
		$r\leq 2$	$r=3$	2.97	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\geq 1$	29.39***	29.68	26.78	31.52	28.70
		$r\leq 1$	$r\geq 2$	9.20	15.41	13.32	17.95	15.66
		$r\leq 2$	$r=3$	2.97	3.76	2.68	8.17	6.50

Note: The critical values reported in all cointegration test tables are taken from Osterwald-Lenum (1992).

Table 4. Cointegration Test: Quarterly Data, 1947:1–87:3 (n = 163)

Variables	Test	H_0	H_1	Test Statistic	Critical Value			
					Trend in DGP		No Trend in DGP	
					95%	90%	95%	90%
I, S	λ_{max}	$r=0$	$r=1$	18.58**	14.06	12.07	14.90	12.91
		$r\leq 1$	$r=2$	2.39	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\geq 1$	20.97**	15.41	13.32	17.95	15.66
		$r\leq 1$	$r=2$	2.39	3.76	2.68	8.17	6.50
I, S, Y	λ_{max}	$r=0$	$r=1$	27.05	20.96	18.59	21.07	18.90
		$r\leq 1$	$r=2$	8.66	14.06	12.07	14.90	12.91
		$r\leq 2$	$r=3$	5.79	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\geq 1$	41.51**	29.68	26.78	31.52	28.70
		$r\leq 1$	$r\geq 2$	14.46	15.41	13.32	17.95	15.66
		$r\leq 2$	$r=3$	5.79	3.76	2.68	8.17	6.50

Table 5. Cointegration Test: Quarterly Data, 1947:1–71:2 ($n = 98$)

Variables	Test	H_0	H_1	Test Statistic	Critical Value			
					Trend in DGP		No Trend in DGP	
					95%	90%	95%	90%
I, S	$\lambda_{\max}$	$r=0$	$r=1$	9.45	14.06	12.07	14.90	12.91
		$r\le 1$	$r=2$	0.82	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\ge 1$	10.28	15.41	13.32	17.95	15.66
		$r\le 1$	$r=2$	0.82	3.76	2.68	8.17	6.50
I, S, Y	$\lambda_{\max}$	$r=0$	$r=1$	32.08	20.96	18.59	21.07	18.90
		$r\le 1$	$r=2$	10.49	14.06	12.07	14.90	12.91
		$r\le 2$	$r=3$	0.02	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\ge 1$	42.60**	29.68	26.78	31.52	28.70
		$r\le 1$	$r\ge 2$	10.52	15.41	13.32	17.95	15.66
		$r\le 2$	$r=3$	0.02	3.76	2.68	8.17	6.50

Table 6. Cointegration Test: Quarterly Data, 1971:3–87:3 ($n = 65$)

Variables	Test	H_0	H_1	Test Statistic	Critical Value			
					Trend in DGP		No Trend in DGP	
					95%	90%	95%	90%
I, S	$\lambda_{\max}$	$r=0$	$r=1$	19.44**	14.06	12.07	14.90	12.91
		$r\le 1$	$r=2$	4.91	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\ge 1$	24.35**	15.41	13.32	17.95	15.66
		$r\le 1$	$r=2$	4.91	3.76	2.68	8.17	6.50
I, S, Y	$\lambda_{\max}$	$r=0$	$r=1$	22.39	20.96	18.59	21.07	18.90
		$r\le 1$	$r=2$	7.24	14.06	12.07	14.90	12.91
		$r\le 2$	$r=3$	4.21	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\ge 1$	33.85**	29.68	26.78	31.52	28.70
		$r\le 1$	$r\ge 2$	11.45	15.41	13.32	17.95	15.66
		$r\le 2$	$r=3$	4.21	3.76	2.68	8.17	6.50
$I/Y, S/Y$	$\lambda_{\max}$	$r=0$	$r=1$	15.09	14.06	12.07	14.90	12.91
		$r\le 1$	$r=2$	1.84	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\ge 1$	16.93**	15.41	13.32	17.95	15.66
		$r\le 1$	$r\ge 2$	1.84	3.76	2.68	8.17	6.50

In the bivariate setting, I find that for the annual data (see Table 3), the null of no cointegration cannot be rejected.

For the quarterly data, *I-S* levels are cointegrated during the 1947:1–87:3 period (see Table 4) as well as during the 1947:1–71:2 period (see Table 5). For the 1971:3–87:3 period, the results are inconclusive because with-trend specification of the test indicates one cointegrating vector but no-trend specification indicates stationarity. When measured in rates (see Table 6), both test statistics indicate *I-S* cointegration with one cointegrating vector.[11]

For the 1980:1–87:3 period, the results support *I-S* cointegration: with no-trend specification, the null of one cointegration vector cannot be rejected. The with-trend cointegration test indicates that the null can be rejected only at 10% significance, but not at 5% significance.

When the variables are measured in rates, I find that during 1971:3–87:3 (see Table 6) and 1980:1–87:3 (see Table 7), both test statistics uniformly reject the null of zero cointegrating vectors in favor of one cointegrating vector. Thus, investment and saving during these periods are cointegrated.

In sum, the bivariate *I-S* cointegration tests are somewhat mixed, although in general they indicate a cointegration if quarterly data are used.

In the trivariate system with *I*, *S*, and *Y*, the results indicate that the three series are cointegrated with one cointegrating vector. This finding holds for all sample periods considered and for both test statistics used (see Tables 4–7). Here we find a cointegration using the annual data also (see Table 3).

Table 7. Cointegration Test: Quarterly Data, 1980:1–87:3 ($n = 31$)

Variables	Test	H_0	H_1	Test Statistic	Critical Value			
					Trend in DGP		No Trend in DGP	
					95%	90%	95%	90%
I, *S*	λ_{max}	$r=0$	$r=1$	17.48**	14.06	12.07	14.90	12.91
		$r\leq 1$	$r=2$	3.08	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\geq 1$	20.57**	15.41	13.32	17.95	15.66
		$r\leq 1$	$r=2$	3.08	3.76	2.68	8.17	6.50
I, *S*, *Y*	λ_{max}	$r=0$	$r=1$	27.23	20.96	18.59	21.07	18.90
		$r\leq 1$	$r=2$	11.84	14.06	12.07	14.90	12.91
		$r\leq 2$	$r=3$	1.80	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\geq 1$	40.88**	29.68	26.78	31.52	28.70
		$r\leq 1$	$r\geq 2$	13.65	15.41	13.32	17.95	15.66
		$r\leq 2$	$r=3$	1.80	3.76	2.68	8.17	6.50
I/Y, *S/Y*	λ_{max}	$r=0$	$r=1$	20.34	14.06	12.07	14.90	12.91
		$r\leq 1$	$r=2$	1.29	3.76	2.68	8.17	6.50
	J_T	$r=0$	$r\geq 1$	21.64**	15.41	13.32	17.95	15.66
		$r\leq 1$	$r\geq 2$	1.29	3.76	2.68	8.17	6.50

This means that the time series of investment and saving are indeed cointegrated, as predicted by the theoretical arguments made in section 2.

The estimated cointegrating vectors and the corresponding adjustment matrices of the cointegration relationships found are reported in Table 8. The long-run coefficient on saving shows remarkable stability with the exception of the annual data, where the estimated coefficient is a little bit higher.[12]

Further, according to the figures reported in Table 8, the homogeneity restrictions seem to be satisfied by the data. For example, in the bivariate regressions, the coefficient on saving is close to 1 whether the regression is run in levels or rates.

Table 8. Cointegrating Vectors and Corresponding Adjustment Matrices

Sample	Cointegrating Vectors					Adjustment Matrixes				
	I	*S*	*Y*	*I/Y*	*S/Y*	*I*	*S*	*Y*	*I/Y*	*S/Y*
1947–87	−1.00	1.19	−0.09			−0.18	−0.33	−0.63		
1947:1–87:3	−1.00	1.00				−0.11	−0.46			
1947:1–87:3	−1.00	1.14	−0.03			−0.27	−0.48	−0.45		
1947:1–71:2	−1.00	1.06	0.14			0.46	0.42	0.21		
1971:3–87:3	−1.00	1.11				−0.80	−1.08			
1971:3–87:3	−1.00	1.14	−0.01			−0.41	−0.70	−0.89		
1971:3–87:3				−1.00	1.06				−0.36	−0.77
1980:1–87:3	−1.00	1.10				−2.87	−3.22			
1980:1–87:3	−1.00	1.05	−0.01			−2.25	−3.39	−2.88		
1980:1–87:3				−1.00	1.05				−2.47	−3.40

Notes: Normalization was carried out by setting the coefficient on investment equal to −1.00. The cointegrating vectors and the adjustment matrices presented here correspond to the cointegration relationships established in Tables 3–7 and are presented in the same order.

Similarly, in trivariate regressions, the sum of the coefficients on *Y* and *S* is close to 1. The speed of adjustment figures reported in the right hand side columns of Table 8, seem rather high. This holds particularly true for the last decade. This suggests that in the US economy, the time series of investment and saving adjust rapidly to their long-run equilibrium levels.

In sum, using the post-war US quarterly and annual data, I find that the time series of investment and saving are cointegrated, which indicates that the U.S. economy is solvent in the sense that it does not violate its dynamic budget constraint. To conclude, therefore, that it is unlikely that *I-S* correlations would provide accurate information on the true degree of international capital mobility.[13]

6. CONCLUSION

Feldstein and Horioka's (1980) finding that saving and investment tend to be correlated in the long run has received significant attention in the literature. This is because Feldstein and Horioka express the view that the long-run *I-S* comovement is an indicator of international capital immobility. If this were true, then the findings reported in this chapter would suggest that capital was not mobile during the 1947–87 period.

As Baxter and Crucini (1993) note, however, most economists disagree with this interpretation. It is difficult to defend this argument for numerous reasons.

First, the restrictions imposed on international capital mobility have been declining over time in the world economy. This is particularly true since early 70s, when many developed, and to a lesser degree developing, countries abolished most capital restrictions.

Second, the increased deregulation and integration of the world financial markets is not compatible with the idea of declining capital mobility. For example, the extreme volatility of exchange rates since the abandonment of the Bretton Woods' system provides persuasive evidence of capital mobility—a large pool of liquid assets are switched in response to anticipation of exchange rate movements.

Third, studies that measure capital mobility directly using various PPP and interest parity conditions, conclude that capital is very mobile and that capital mobility has been increasing over time. For example, Hutchison and Singh (1993) examine real interest rate differential between the U.S. and Japan and find that capital mobility is very high. Popper (1990) uses interest and currency arbitrage conditions along with financial asset returns and finds that capital is as mobile in the long- as in the short-run.

This chapter claims that there is nothing mysterious in the *I-S* comovement. Since the neoclassical growth theory predicts that in the steady state investment and saving should be proportional to output and therefore would grow at the same rate, it would be surprising if we did not find a high long-run *I-S* correlation. The modern optimization-based dynamic model of open economy also predicts that investment and saving would be correlated in the long run regardless of the extent of capital mobility, unless the economy violates its dynamic budget constraint. Therefore, a test of *I-S* cointegration is merely a test of country's economic solvency. To conclude, therefore, the observed long-run *I-S* correlation cannot be useful for measuring the extent of international capital mobility.

As additional evidence, it should be noted that if Feldstein and Horioka line of argument were valid, then the huge capital inflow to the U.S. during the first term of the Reagan administration should have diminished the long-run *I-S* correlation in the early 80s. The findings reported here, however, do not support this view.

NOTES

1. See, for example, Sinn (1992), Ghosh (1995), Coakley, Kulasi, and Smith (1996), Sachsida and Caetano (2000) and the references cited therein. More recent studies include Tsoukis and Alyousha (2001), Alyousha and Tsoukis (in this volume), and Fountas and Tsoukis (2000), who study a sample of seven industrialized economies.
2. See, for example, Bayoumi (1990), Sachs (1981), Obstfeld (1986), Frankel (1991), Levy (1995), Frankel and MacArthur (1988), Popper (1990), and Baxter and Crucini (1993).
3. I shall mention that the focus of this chapter is the long-run capital mobility. Short-run capital mobility is less controversial. See, for example, Feldstein (1983) and Levy (2000, 2001).
4. Virtually all other macro models, with or without open capital markets, make similar predictions on the long-run investment-saving comovement.
5. For example, Barro, Mankiw, and Sala-I-Martin (1995) construct an open economy version of the neoclassical growth model with this conclusion.
6. A similar problem arises in the estimation of cointegrating regressions using Engle-Granger two-step method, where the residuals' ADF unit root test statistic distribution depends on the true value of the intercept term. As MacKinnon (1991) notes, all tables assume that $\alpha_1 = 0$, and, therefore, may be quite misleading if this is not the case.
7. The reason for the common use of investment and saving rates is to avoid the difficulties the presence of integrated variables create in traditional regression analysis. It turns out, however, that modeling the time series of investment and saving as rates may not be sufficient to make them stationary. See footnote 10 below.
8. Miller (1988) considers only investment and saving rates, Otto and Wirjanto (1989) only consider levels, and Gulley (1992) considers both levels and rates.
9. To conserve degrees of freedom, I do not divide the annual data into sub-periods because the cointegration test I use employs a maximum likelihood procedure based on error correction representation of the VAR formed by the variables considered.
10. It may seem puzzling that investment and saving are $I(1)$ in levels as well as rates. See Levy (2000, p. 115, footnote 13) for a possible explanation.
11. Because of the possibility that investment and saving rates may contain no deterministic trend, the data were also tested for cointegration with the restriction $\mu = \mathbf{0}$, where μ is the vector of constants in the VAR. The result is identical. That is, the time series of investment and saving were found to be cointegrated with one cointegration vector.
12. Sinn (1992) also finds that the coefficient is higher for lower frequency data.
13. I have also tested for *I-S* cointegration in rates under the assumption that the series don't contain a deterministic trend. The results are the same. It should also be noted that the investment-saving relationships studied here were also estimated recursively and the results indicate significant parameter constancy.

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