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Unit root tests of the current account balance: implications for international capital mobility

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UNIT ROOT TESTS OF THE CURRENT ACCOUNT
BALANCE: IMPLICATIONS FOR
INTERNATIONAL CAPITAL MOBILITY

by

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Abstract

Unit Root Tests Of The Current Account Balance: Implications For International Capital Mobility

by

Erich Gundlach and Stefan Sinn

This paper assesses the extent of international capital mobility in a time series context. It explores the possibility that the current account balance of different OECD-countries contains a unit root. It is shown that if the ratio of the current account balance to GDP is found to be integrated of the order of one, the country is likely to be part of the world capital market. The results for the whole period 1950-1988 indicate that the current account balance of at least Germany, Japan, and the United States contains a unit root. Considering the subperiods before and after 1972 it is shown that international capital mobility increased after the breakdown of the Bretton Woods System.

JEL Classification: F 21, F 32

Key words: International capital mobility, current account adjustment, unit root tests

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Unit Root Tests of the Current Account
Balance: Implications for
International Capital Mobility*

by

Erich Gundlach and Stefan Sinn

I. Introduction

Is capital mobility among industrialized countries high or low? To the businessman involved in financial markets the answer seems obvious: yes, international capital mobility is high and growing more so every day. The suggestion that capital mobility among industrialized countries is low would at first sight appear incredulous to the daily participants in stock exchanges and foreign exchange trading. Nevertheless, this is the conclusion drawn from a number of recent studies attempting to empirically assess the extent of international capital mobility using correlations of saving and investment rates. This approach to the measurement of international capital mobility was first suggested by Feldstein, Horioka (1980). It is based on the idea that with international capital mobility it is unlikely that at any point in time high saving countries are also countries with high investment rates. If, on the other hand, the economy is completely closed to capital movements, changes in savings must be accompanied by changes in investment. Therefore, if a regression of the investment rate on the saving rate yields a parameter value which is statistically not different from one, this would suggest international capital immobility.

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Empirical evidence put forward in Feldstein, Horioka (1980), Feldstein (1983), and Feldstein, Bacchetta (1989) shows that for the group of OECD countries this parameter value is close to one in the 1960s and (surprisingly so) in the 1970s. In the 1980s the parameter value is somewhat lower than in the previous two decades. These results seem to indicate that international capital mobility is far from perfect. Economists have responded to this puzzling finding by criticizing the method employed by Feldstein, Horioka (1980) on statistical and theoretical grounds.¹ We believe that these criticisms have not been sufficient to challenge the gist of Feldstein and Horioka's findings and that the subject merits further research. In particular, we argue that tests of saving and investment correlations based on the cross-section approach usually employed in the literature are plagued by a number of problems. These problems lead us to employ a time series approach for testing whether or not a country has been linked to the international capital market.

Our testing approach explores the possibility that the current account balance of different OECD-countries contains a unit root. We show that if the ratio of the current account balance to GDP is found to be integrated of the order of one, the existence of a stable long-term relationship between the saving and investment rates of that country is unlikely. Therefore any inferences based on such a specification may be regarded as spurious.

It cannot be concluded, however, that a country is shut off from the international capital market if its current account balance is found to be $I(0)$. A number of theoretical studies

¹ See the summaries of these critiques in Dooley, Frankel, Mathieson (1987) and Sinn (1991).

suggest that over time both saving and investment rates are influenced by the same exogenous variables.² In that case saving and investment rates could be cointegrated and the current account balance would have to be $I(0)$ even if the country is linked to the international capital market. Up to now conclusive evidence in favor of these models is missing. Therefore, our tests should be seen as a preliminary assessment of international capital mobility in a time series context.

II. Outline of the test procedure.

In line with Feldstein, Horioka (1980) most authors have performed saving and investment correlations on a cross-section of OECD-countries. The regression equation estimated is of the general form

$$[1] \quad I/Y = \alpha + \beta S/Y + e$$

where I/Y is the ratio of investment to GDP and S/Y the ratio of saving to GDP and e is an error term where the errors are assumed to be independent and identically distributed with a mean of 0 and a variance of σ_e^2 (iid with $(0, \sigma_e^2)$).

There are at least three reasons why a time series estimation of equation [1] might be of interest.³ First, a time series approach more accurately reflects the theoretical set up of most models. In particular, models of small open economies typically predict how much of an increase in saving ends up as

² Compare for example Obstfeld (1986), Murphy (1986), Cardia (1988), Tesar (1988), Baxter, Crucini (1990), Engel, Kletzer (1990), Wong (1990), Leachman (1991). Most authors identify movements of the business cycle as an exogenous source that might cause saving and investment rates to move in the same direction.

³ Compare also Obstfeld (1986).

domestic investment. None, theory tells us, if capital is mobile, and all if it is not. Adopting this idea to a cross-section encounters the conceptual difficulty that one is measuring events at one particular point in time in different countries.⁴ Theory only suggests that with international capital mobility it is unlikely that at any point in time high saving countries are also countries with high investment rates. But it does not rule out that one might observe this in spite of capital mobility.⁵

Second, the average one obtains over a cross-section of countries might be the result of divergent individual observations. Countries may contribute to a high β -coefficient either because they have imposed capital controls or because they are large.⁶ Thus the results obtained in the cross-section analysis by Feldstein, Horioka (1980) and subsequent work are difficult to interpret since they are based on a sample that includes small and large countries.

Third, the recently introduced concept of cointegration (Granger, 1981; Engle, Granger, 1987) suggests that it is possible to estimate long run equilibrium relationships from time series data in levels, even though the data may not exhibit stationarity. That is, a differencing of the data to avoid the "spurious regression" problem is not necessary if certain empirically testable conditions hold. Therefore, the cointegra-

⁴ Nevertheless Feldstein, Bacchetta (1989) refer to the coefficient β in equation [1] as the "saving retention coefficient", effectively treating another country in the cross-section as the same country at another point in time.

⁵ Cross-section studies have tried to incorporate time in their results by basing them on long-term averages. However, this biases their results towards the acceptance of the hypothesis of capital immobility as shown in Sinn (1991).

⁶ This point is made by Murphy (1984); Obstfeld (1986) shows that large countries tend to have a higher correlation between saving and investment rates.

tion approach allows to retain the long run information inherent in the data, which is lost by the differencing procedure. Since the capital mobility controversy focusses on long run equilibrium relationships, the cointegration approach seems to offer some practical guidelines for the conduct of the time series analysis of this issue.⁷

We are not the first to explore this issue in a time series context. Obstfeld (1986) looked at the correlation between quarterly first differences in savings and investment rates for seven industrialized countries for the period 1959 to 1984. As noted above, his findings show that the correlation coefficient increases with country size and that (for a reduced sample of six countries) it is lower during the period after the breakdown of the Bretton Woods System for all but one country. Miller (1988) finds that U.S. quarterly saving and investments rates (1946 I to 1987 III) are cointegrated during the area of fixed exchange rates but not so under flexible exchange rates. Leachman (1991) tests for cointegration between the saving and investment rates of 24 OECD-countries.⁸ She rejects cointegration for all 24 countries and concludes, contrary to Feldstein, Horioka (1980), that their capital markets are not closed. Leachman's conclusion is, however, open to question, especially if one compares her results with the critical values for cointegration tests tabulated in Philipps, Ouliaris (1990).

⁷ Feldstein, Horioka (1980, p. 323) stress the fact that they view equation [1] "... as a long run relation in which intercountry differences in saving rates reflect basic structural differences among countries."

⁸ Leachman (1991) uses net saving and investment rates. Feldstein, Horioka (1980) argue that gross rates are preferable because it is gross flows that are mobile between countries and because the measurement of depreciation is open to doubt.

Our test procedure simplifies the approach used by these authors which is based on testing for cointegration between saving and investment rates. If saving and investment are cointegrated, equation [1] is well-specified and yields an unbiased estimate of β . If β equals 1 (if it is not statistically different from 1) one would reject the hypothesis of international capital mobility. Since the current account balance equals the difference between saving and investment equation [1] can be rewritten as

$$[2] \quad CA/Y = -\alpha + [1-\beta] S/Y - e$$

where CA/Y is the current account surplus divided by GDP.

Now suppose that the current account balance is found to be $I(1)$. If we stick to the implicit Feldstein-Horioka assumption that the error term e is iid with $(0, \sigma_e^2)$, then the current account balance can exhibit $I(1)$ behavior only in the case where β does not equal 1. Here the $I(1)$ result would indicate different reactions of the saving and the investment rate to shocks. That is, in this case the saving and the investment rate can be interpreted as following independent random walks. If β indeed equals 1 and the error term is stationary, then the country under consideration must have an $I(0)$ current account and its saving and investment rate will not show different reactions to shocks. Therefore, a country is necessarily linked to the international capital market if its current account balance is found to be $I(1)$.

Of course there is the theoretical possibility that β equals 1 and e is not stationary. If this were true, one would have perfect capital immobility although our test procedure would indicate a unit root in the current account balance. As an example consider the case of a developing country which is closed to the international capital market, but irregularly receives international aid in goods or financial assistance.

Our test procedure would falsely identify such a country as being part of the world capital market.

We admit that our approach to testing for international capital mobility may be misleading for countries which are completely closed to the international capital market but occasionally receive or transfer resources. We want to stress that this problem is not only present in our approach but also in the cointegration approach taken e.g. by Leachman (1991). But we do not think that this possible bias towards the acceptance of a unit root (towards the rejection of international capital immobility) is of practical relevance for a sample of OECD-countries, which we test. Therefore, we suggest that testing whether the current account balance expressed as a ratio to GDP contains a unit root is a meaningful test for international capital mobility for a sample of OECD-countries, and that this approach is equivalent to but simpler than testing for cointegration between saving and investment rates.

III. Unit root tests of the current account balance.

Unit root tests are known for their low power to discriminate against close local alternatives, especially in situations with highly trending data (Cochrane (1991)). As a consequence, these tests should not be expected to reveal the "true" data generating process. But testing whether the current account balance contains a unit root avoids the low power problem to some extent, because we are only interested in establishing whether the current account balance is stationary ($I(0)$) or not. If we find that the current account balance is not $I(0)$, it is irrelevant for our argument whether this property is due to a "true" simple random walk, to a "true" trend-stationary-process or to a "true" difference-stationary-process. In all these cases, there can be no long run stable relationship between the saving rate and the investment rate, and any inferences drawn from such a specification can be regarded as

spurious.⁹ Further, if there is no stable long run relationship between the saving rate and the investment rate, the country under consideration is linked to the international capital market by definition.

We consider two basic parameterisations to model the data generating process of the current account balance:

$$[6] \quad y_t = \beta_1 y_{t-1} + u_t$$

$$[7] \quad y_t = c + \beta_2 y_{t-1} + u_t$$

where y is the time series of the current account balance of the country under consideration, c is a regression constant, β is a constant parameter and u is an error term, while the subscript t refers to the time period. A priori, we expect that the current account balance is one of the few economic time series which does not exhibit a trend. Therefore, it is not necessary to include a trend component. The first parameterisation (equation [6]) does not allow for a mean current account balance different from zero. We regard this as the most realistic scenario. However, to avoid a bias towards the non-rejection of a unit root in our testing procedures, we allow for a non-zero mean current account balance in the second parameterisation (equation [7]). This parameterisation may be relevant for countries which exhibit a permanent positive or negative current account balance.¹⁰ In the remainder of the paper, we refer to the first and second parameterisation as model 1 and model 2.

⁹ This argument does not hold if one allows the saving rate and the investment rate to follow different deterministic trends, which means that the current account balance follows a deterministic trend, too. However, we are not aware of a theory which could explain a deterministic trend in the current account balance.

¹⁰ This paper does not attempt to explain why countries should have a persistent non-zero mean current account balance.

To test for a unit root in models (1) and (2) we use the procedures developed by Dickey, Fuller (1979), Said, Dickey (1984), Phillips (1987), and Phillips, Perron (1988). Critical values for these unit root tests are tabulated in Fuller (1976, p. 373). The first two procedures are labelled Dickey-Fuller (DF) and Augmented-Dickey-Fuller (ADF) tests, while the latter two are usually referred to as the Phillips-Perron $Z(t_\alpha)$ -statistics.

The DF and the ADF test equations for models 1 and 2 are OLS-regressions which read as follows:

$$[8] \quad \text{DF1} : \Delta y_t = \alpha y_{t-1} + e_t$$

$$[9] \quad \text{ADF1} : \Delta y_t = \alpha y_{t-1} + \sum_{i=1}^1 \gamma_i \Delta y_{t-1} + e_t$$

$$[10] \quad \text{DF2} : \Delta y_t = c + \alpha y_{t-1} + e_t$$

$$[11] \quad \text{ADF2} : \Delta y_t = c + \alpha y_{t-1} + \sum_{i=1}^1 \gamma_i \Delta y_{t-1} + e_t$$

with e_t iid $(0, \sigma_e^2)$. If the null-hypothesis $H_0 : \alpha = 0$ cannot be rejected, the time series under consideration is supposed to contain a unit root.

The Phillips-Perron $Z(t_\alpha)$ -statistic is intended to correct the conventional regression statistic from, say, a DF-test for the effects of heteroscedasticity and autocorrelation in the error terms of the models (1) and (2) described in equation [6] and [7]. The $Z(t_\alpha)$ -statistics for models (1) and (2) therefore provide a very general test for the presence of a unit root. They are defined by equations [12] and [13]:

$$[12] \quad Z(t_\alpha)1 = \left[\sum_1^T y_t^2 - 1 \right]^{1/2} [\beta_1 - 1] / s_{T1} - 0.5 [s_{T1}^2 - s_u^2] \left[\sum_1^T y_{t-1}^2 \right]^{1/2} \left[T^{-2} \sum_1^T y_{t-1}^2 \right]^{-1}$$

where
$$s_u^2 = T^{-1} \sum_1^T u_t^2$$

and u_t is the residual from an OLS-regression of equation [6]

and
$$s_{T1}^2 = T^{-1} \sum_1^T u_t^2 + 2T^{-1} \sum_{\tau=1}^1 w_{\tau 1} \sum_{t=\tau+1}^T u_t u_{t-\tau}$$

where $w_{\tau 1} = 1 - \tau/[l+1]$.

[13]
$$Z(t_\alpha)^2 = [s_u / s_{T1}] [\beta_2 - 1] / s_{\beta_2} - \lambda' s_{T1} / \bar{m}_{yy}^{1/2}$$

where $\lambda = 0.5 [s_{T1}^2 - s_u^2]$

and $\lambda' = \lambda / s_{T1}^2$

and $\bar{m}_{yy} = T^{-2} \sum (y_t - \bar{y})^2$

and s_{β_2} is the standard error of β_2 in an OLS-regression of

equation [7], s_u and s_{T1} are defined as before, and here u_t is the residual from an OLS-regression of equation [7].

The time series of the current account balance are calculated from the Summers, Heston (1991) data set using the current international dollar series for the years 1950 to 1988. That is, we calculate the current account balance for each country as 100 minus the consumption share of GDP (column cc of the Summers and Heston data set) minus the investment share of GDP (column ci) minus the government share of GDP (column cg). The results of our unit root tests of the current account balance are listed in Tables 1-6, beginning with the DF- and the $Z(t_\alpha)$ -tests of models (1) and (2) for the period 1950-88 (see Tables 1-4) and followed by tests for the subperiods 1950-72 and 1973-88 (see Tables 5 and 6). These tables present the estimated t-statistics of the coefficient α of the specified test equation for alternative models and lag structures.

We begin our analysis with the simple DF test for model 1 (Table 1). The first column gives the t-statistic for the estimated coefficient α of an OLS-regression of equation [8]; the next two columns give the t-statistics for the estimated coefficient of an OLS-regression of equation [9] with lag lengths $l = 1$ and $l = 2$. Of course, the lag length is a purely empirical question. For annual data spanning a period of approximately 40 years as in the present case one could argue that very often the inclusion of one lag suffices to produce white noise residuals of the test equation. However, it is important to note that the lag length chosen may seriously affect the results. For this reason Campbell, Perron (1991) suggest data based procedures to estimate the "correct" lag structure. Since such procedures in turn may have their drawbacks, here we simply use alternative lag structures to check the stability of our results.

Whether the current account balance of a country contains a unit root can be assessed by comparing the t-statistics of Table 1 with the critical values provided by Fuller (1976). When interpreted conservatively these critical values indicate that one can reject the null-hypothesis of a unit root if the estimated t-statistic is smaller than -1.61, the approximate critical value for a 10 percent level of statistical significance. With a less conservative interpretation, one rejects the null-hypothesis of a unit root if the estimated t-statistic is smaller than -1.91, the approximate critical value for a 5 percent level of significance. Given that unit root tests have low power in discriminating against close alternatives, we suggest that one should use the 10 percent critical value in order to avoid a bias in judgement in favour of the unit root hypothesis.

For some countries, the results are somewhat unclear. For example, consider the case of the United States. Without a lag in the test equation, one accepts the unit root hypothesis. Including one lag leads to a rejection of the unit root

Table 1 - Dickey-Fuller tests of the current account balance
1950-1988, Model 1^{a,b}

| | (1) Equation [8] no lag | (2) Equation [9] one lag | (3) Equation [9] two lags |
|----------------|-------------------------------|--------------------------------|---------------------------------|
| Australia | -6.93* | -4.25* | -1.92* |
| Austria | -4.01* | -5.57* | -2.37* |
| Belgium | -2.52* | -1.67* | -1.57 |
| Canada | -2.50* | -1.85* | -1.95* |
| Denmark | -2.32* | -1.48 | -1.02 |
| Finland | -3.86* | -4.51* | -3.17* |
| France | -4.39* | -4.08* | -3.75* |
| Germany | 0.18 | -0.20 | 0.50 |
| Greece | -1.27 | -0.99 | -0.62 |
| Iceland | -3.84* | -4.01* | -2.55* |
| Ireland | -1.12 | -1.56 | -1.07 |
| Italy | -2.79* | -2.59* | -2.06* |
| Japan | -1.37 | -1.01 | -1.01 |
| Netherlands | -1.98* | -1.98* | -0.93 |
| New Zealand | -4.73* | -3.64* | -2.46* |
| Norway | -2.47* | -3.05* | -2.88* |
| Portugal | -0.87 | -1.00 | -0.88 |
| Spain | -2.67* | -2.93* | -2.83* |
| Sweden | -3.06* | -2.46* | -2.25* |
| Switzerland | -3.88* | -4.50* | -3.66* |
| Turkey | -1.77* | -1.37 | -1.15 |
| United Kingdom | -2.27* | -2.58* | -2.51* |
| United States | -1.38 | -2.04* | -1.53 |

^a Without constant.

^b Critical values: -1.95 (5 p.c.); -1.61 (10 p.c.); *: null-hypothesis of a unit root is rejected.

Source: Summers, Heston (1991).

hypothesis, and including two lags reestablishes the first result. For model (2) (Table 2) we find a similar instability for some countries. Here one rejects the null-hypothesis of a unit root if the estimated t -statistic is smaller than -2.61 , the approximate critical value for a 10 percent level of statistical significance.

A clearer picture emerges if one uses the $Z(t_\alpha)$ -test for model (1) (Table 3). We use arbitrarily chosen lag lengths of 1, 2, 3, and 6 years. Our calculations show that the lag length seems to have no important impact on the stability of the test results.¹¹ Again the case of the United States is somewhat unclear since the t -statistics for the lag lengths $l=1$ and $l=2$ are quite close to the critical value. Nevertheless, one may interpret the evidence as pointing to the acceptance of the unit root hypothesis. Given the empirical evidence produced by this test, Germany, Greece, Ireland, Japan, Portugal, the United States, and possibly Turkey seem to have a non-stationary current account balance. With the exception of Turkey and Greece this result is confirmed by the $Z(t_\alpha)$ -test of model (2) (Table 4). And note that this result is also supported by the DF tests (Tables 1 and 2).

Broadly speaking, the picture that our test procedures reveal is that for some major countries, namely Germany, Japan, and the United States, there is empirical evidence that their current account balance contains a unit root. That is, there seems to be no long run stable relationship between the saving and the investment rates of these countries which in turn implies that they are necessarily a part of the international capital market. Together with the results for some smaller countries where the evidence is not as clear cut if one uses conservative levels of statistical significance, this contradicts the

¹¹ This finding is not an implication of the test statistic. See, for instance, Perron (1990).

Table 2 - Dickey-Fuller tests of the current account balance
1950-1988, Model 2^{a,b}

| | Equation [10] no lag | Equation [11] one lag | Equation [11] two lags |
|----------------|-------------------------|--------------------------|---------------------------|
| Australia | -7.54* | -4.50* | -2.36 |
| Austria | -3.99* | -5.65* | -2.44 |
| Belgium | -2.51 | -1.66 | -1.56 |
| Canada | -2.58 | -1.97 | -1.97 |
| Denmark | -2.23 | -1.30 | -0.82 |
| Finland | -3.81* | -4.42* | -3.12* |
| France | -4.60* | -4.47* | -4.36* |
| Germany | -1.44 | -1.69 | -0.96 |
| Greece | -2.65* | -2.35 | -1.99 |
| Iceland | -4.20* | -4.55* | -2.80* |
| Ireland | -0.52 | -0.96 | -0.84 |
| Italy | -3.75* | -3.72* | -2.90* |
| Japan | -1.77 | -1.40 | -0.15 |
| Netherlands | -2.76* | -2.72* | -1.39 |
| New Zealand | -4.92* | -3.77* | -2.62* |
| Norway | -2.44 | -3.06* | -2.87* |
| Portugal | -1.97 | -2.53 | -2.16 |
| Spain | -3.17* | -3.73* | -3.94* |
| Sweden | -3.12* | -2.49 | -2.29 |
| Switzerland | -3.84* | -4.44* | -3.64* |
| Turkey | -2.84* | -2.24 | -1.61 |
| United Kingdom | -2.41 | -2.65* | -2.65* |
| United States | -1.17 | -1.92 | -1.41 |

^a With constant.

^b Critical values: -2.95 (5 p.c.); - 2.61 (10 p.c.); *: null-hypothesis of a unit root is rejected.

Source: See Table 1.

Table 3 - $Z(t_\alpha)$ -tests^a of the current account balance 1950-1988
Model 1^{b,c}

| | Lag length | | | |
|----------------|------------|--------|--------|--------|
| | 1 = 1 | 1 = 2 | 1 = 3 | 1 = 6 |
| Australia | -6.97* | -6.95* | -6.95* | -7.03* |
| Austria | -4.08* | -3.96* | -3.94* | -4.01* |
| Belgium | -2.65* | -2.75* | -2.86* | -2.92* |
| Canada | -2.42* | -2.47* | -2.56* | -2.61* |
| Denmark | -2.04* | -2.11* | -2.87* | -2.50* |
| Finland | -3.95* | -3.90* | -3.80* | -3.75* |
| France | -4.49* | -4.67* | -4.38* | -4.30* |
| Germany | -0.01 | 0.11 | 0.23 | 0.20 |
| Greece | -1.29 | -1.28 | -1.29 | -1.30 |
| Iceland | -3.94* | -3.80* | -3.80* | -3.71* |
| Ireland | -1.22 | -1.28 | -1.30 | -1.40 |
| Italy | -2.84* | -2.74* | -2.72* | -2.74* |
| Japan | -1.36 | -1.20 | -1.21 | -1.33 |
| Netherlands | -2.18* | -2.19* | -2.21* | -2.11* |
| New Zealand | -4.84* | -4.81* | -4.83* | -4.90* |
| Norway | -2.72* | -2.76* | -2.71* | -2.52* |
| Portugal | -0.99 | -0.98 | -0.91 | -0.71 |
| Spain | -2.83* | -2.85* | -2.77* | -2.49* |
| Sweden | -3.05* | -3.10* | -3.18* | -3.12* |
| Switzerland | -4.04* | -3.96* | -3.87* | -3.62* |
| Turkey | -1.66 | -1.57 | -1.58 | -1.69* |
| United Kingdom | -2.39* | -2.47* | -2.39* | -2.27* |
| United States | -1.54 | -1.55 | -1.51 | -1.36 |

^a See equation [12]

^b Without constant.

^c Critical values: -1.95 (5 p.c.); -1.61 (10 p.c.); *: null-hypothesis of a unit root is rejected.

Source: See Table 1.

Table 4 - $Z(t_\alpha)$ -tests^a of the current account balance 1950-1988
Model 2^{b,c}

| | Lag length | | | |
|----------------|------------|--------|--------|--------|
| | l = 1 | l = 2 | l = 3 | l = 6 |
| Australia | -7.43* | -7.41* | -7.37* | -7.33* |
| Austria | -3.99* | -3.86* | -3.84* | -3.90* |
| Belgium | -2.59 | -2.67* | -2.77* | -2.82* |
| Canada | -2.48 | -2.53 | -2.63* | -2.66* |
| Denmark | -2.09 | -2.07 | -2.28 | -2.42 |
| Finland | -3.86* | -3.80* | -3.69* | -3.64* |
| France | -4.65* | -4.60* | -4.50* | -4.41* |
| Germany | -1.69 | -1.68 | -1.64 | -1.73 |
| Greece | -2.69* | -2.73* | -2.80* | -3.02* |
| Iceland | -4.26* | -4.12* | -4.20* | -3.94* |
| Ireland | -0.64 | -0.74 | -0.75 | -0.92 |
| Italy | -3.83* | -3.76* | -3.71* | -3.58* |
| Japan | -1.77 | -1.66 | -1.67 | -1.76 |
| Netherlands | -2.88* | -2.90* | -2.92* | -2.81* |
| New Zealand | -4.95* | -4.92* | -4.92* | -4.93* |
| Norway | -2.67* | -2.71* | -2.66* | -2.46 |
| Portugal | -2.14 | -2.16 | -2.10 | -1.90 |
| Spain | -3.30* | -3.31* | -3.20* | -2.78 |
| Sweden | -3.08* | -3.13* | -3.20* | -3.14* |
| Switzerland | -3.95* | -3.86* | -3.77* | -3.49* |
| Turkey | -2.80* | -2.77* | -2.82* | -3.00* |
| United Kingdom | -2.50 | -2.58 | -2.51 | -2.39 |
| United States | -1.35 | -1.37 | -1.32 | -1.20 |

^a See equation [13].

^b With constant.

^c Critical values: -2.95 (5 p.c.); -2.61 (10 p.c.); *: null-hypothesis of a unit root is rejected.

Source: See Table 1.

Feldstein-Horioka conclusion that within the group of OECD-countries, international capital mobility is surprisingly low.

This interpretation is reinforced if one analyzes the two sub-periods 1950-72 and 1973-88. Our motivation for splitting the sample into these periods is that we expect that the degree of international capital mobility as measured by the Feldstein-Horioka test increased after the breakdown of the Bretton Woods System of fixed exchanged rates in 1972. While it is indisputable that net international capital flows occurred under the Bretton Woods System, we expect that they did not respond as freely to shocks in investment opportunities or saving behavior as they did under the system of flexible exchange rates. Instead, government interventions that were motivated by current account targets - a crucial element of the Bretton Woods System - are likely to have lowered the mobility of net international capital flows. Therefore, we expect that an analysis of the whole period 1950-88 may bias the results towards the acceptance of the hypothesis of low international capital mobility. By splitting the sample, we expect to find an increase in the international mobility of capital as measured by our test after 1972 because current account targets lose their importance under a system of flexible exchange rates.

Table 5 presents our test results for the subsamples 1950-72 and 1973-88 using the two augmented Dickey-Fuller tests. Here we only choose test equations with a lag length of one year. A further extension of the lag length is restricted by the relatively small sample sizes, so this is the most general model one can use. The overall impression one gets from these Dickey-Fuller tests is that the computed t-statistics are generally much lower for the post Bretton Woods period. There is a substantially higher number of countries which exhibit an $I(0)$ current account balance in the period 1950-72 compared to the period 1973-88: According to the DF-tests we find 19 countries with an $I(0)$ current account balance out of a sample of 23 in the former period, whereas in the latter period we find

Table 5 - Augmented Dickey-Fuller tests^a of the current account balance, 1950-1972 and 1973-1988

| | 1950 - 1972 | | 1973 - 1988 | |
|----------------|------------------------------|-------------------------------|------------------------------|-------------------------------|
| | Equation [9] ^b | Equation [11] ^c | Equation [9] ^b | Equation [11] ^c |
| Australia | -6.14* | -5.87* | -1.74 | -2.58 |
| Austria | -4.78* | -6.22* | -3.07* | -3.11* |
| Belgium | -0.89 | -0.79 | -1.38 | -1.34 |
| Canada | -1.30 | -1.23 | -1.22 | -1.57 |
| Denmark | -1.96* | -2.32 | -0.73 | -0.49 |
| Finland | -4.35* | -4.15* | -2.99* | -2.88* |
| France | -3.53* | -5.30* | -2.64* | -2.53 |
| Germany | -0.64 | -2.80* | -0.01 | -1.39 |
| Greece | -0.90 | -2.04 | -0.43 | -2.13 |
| Iceland | -3.34* | -4.52* | -2.34* | -2.37 |
| Ireland | -2.38* | -6.91* | -0.74 | -0.23 |
| Italy | -2.02* | -2.39* | -1.63* | -3.25* |
| Japan | -1.59 | -1.65 | -0.55 | -1.07 |
| Netherlands | -5.01* | -4.86* | -0.60 | -1.83 |
| New Zealand | -5.35* | -5.66* | -1.83* | -2.48 |
| Norway | -2.30* | -3.92* | -1.96* | -1.88 |
| Portugal | -0.18 | -2.72* | -0.75 | -2.81* |
| Spain | -1.80* | -2.44 | -1.95* | -2.45 |
| Sweden | -2.63* | -2.56 | -1.51 | -1.56 |
| Switzerland | -3.50* | -3.52* | -2.86* | -2.77* |
| Turkey | -1.12 | -3.65* | -0.86 | -1.31 |
| United Kingdom | -3.76* | -3.67* | -1.34 | -1.30 |
| United States | -1.60 | -3.43* | -1.43 | -1.35 |

^a Lag length l=1.

^b Critical values: -1.95 (5 p.c.); -1.61 (10 p.c.).

^c Critical values: - 2.95 (5 p.c.); -2.61 (10 p.c.).

* Unit root hypothesis is rejected at 10 p.c. level.

Source: See Table 1.

10 countries with an $I(0)$ current account balance. As before, the $Z(t_\alpha)$ -tests show the same result (Table 6): The t -statistics are substantially lower for the post Bretton Woods era, and the difference in the number of countries with an $I(0)$ current account balance is even greater.

Taken together, we interpret the relative stability of our results across different test procedures as an indication of an increase in international capital mobility since the breakdown of the Bretton Woods System. However, we warn the reader not to overinterpret the findings, since the relatively small number of observations for the two subperiods can introduce a bias towards the non-rejection of the unit root hypothesis. In particular, we suggest that our results should not be interpreted as having identified the "true" data generating process of the current account balance of single countries. It is not possible to extract this information from the present small samples. But we think that the different pattern for the two subsamples and its stability when submitted to different test procedures does contain information, namely that the current account balance shows more unpredictable variability since the breakdown of the Bretton Woods System. That is, for that period our null-hypothesis of a unit root in the current account balance can not as easily be rejected as for the former period. As we have set out in section II, this finding can be interpreted as an indication of an increase in the international mobility of capital.

IV. Conclusions

This paper develops a new approach to testing whether a country is linked to the international capital market. We show that if a country's current account balance expressed as a ratio to GDP contains a unit root, then there is no long run stable relationship between its saving and investment rates and the country is linked to the international capital market. Our results for the whole period from 1950 to 1988 indicate that at least Germany, Japan, and the United States are part of the inter-

Table 6 - $Z(t_{\alpha})$ -tests of the current account balance^a 1950-1972 and 1973-88

| | 1950-72 | | 1973-88 | |
|----------------|----------------------|----------------------|----------------------|----------------------|
| | Model 1 ^b | Model 2 ^c | Model 1 ^b | Model 2 ^c |
| Australia | -7.96* | -7.74* | -2.11* | -2.97* |
| Austria | -2.90* | -2.91* | -2.88* | -2.87* |
| Belgium | -3.35* | -3.13* | -1.21 | -1.23 |
| Canada | -2.15* | -2.04 | -1.38 | -1.75 |
| Denmark | -2.86* | -3.14* | -1.13 | -0.94 |
| Finland | -3.94* | -3.77* | -2.19* | -2.12 |
| France | -2.84* | -3.22* | -3.54* | -3.45* |
| Germany | -0.70 | -3.96* | 0.27 | 0.97 |
| Greece | -1.34 | -2.25 | -0.56 | -2.65* |
| Iceland | -2.86* | -3.37* | -2.57* | -2.59 |
| Ireland | -1.51 | -3.21* | -0.60 | -0.07 |
| Italy | -1.89* | -2.23 | -1.92* | -3.03* |
| Japan | -1.97* | -2.04 | -0.72 | -1.20 |
| Netherlands | -4.00* | -3.78* | -0.52 | -1.71 |
| New Zealand | -6.13* | -5.94* | -2.36* | -2.88* |
| Norway | -3.12* | -3.84* | -1.61 | -1.56 |
| Portugal | -0.82 | -3.93* | -0.59 | -1.94 |
| Spain | -1.89* | -2.20 | -1.81* | -2.20 |
| Sweden | -2.42* | -2.34 | -2.05* | -2.03 |
| Switzerland | -2.58* | -2.61 | -2.69* | -2.60 |
| Turkey | -1.68* | -4.47* | -0.96 | -1.54 |
| United Kingdom | -2.93* | -3.31* | -1.10 | -1.08 |
| United States | -1.23 | -2.66* | -1.03 | -0.93 |

^a Lag length l=1.

^b Without constant, see equation [12]; critical values: -1.95 (5 p.c.),
-1.61 (10 p.c.).

^c With constant, see equation [13]; critical values: -2.95 (5 p.c.);
-2.61 (10 p.c.).

* Unit root hypothesis is rejected at 10 p.c. level.

Source: See Table 1.

national capital market. Considering the subperiods before and after the breakdown of the Bretton Woods System, our tests show that there is a marked increase in the extent of international capital mobility after 1972. We caution the reader not to construe this result as having identified the "true" extent of international capital mobility before and after the breakdown of the Bretton Woods System. In particular, our results might be rather fragile due to the low number of observations in the two subsamples. Nevertheless, the fact that the same pattern of low and high capital mobility before and after 1972 emerged under different test procedures convinces us that our results are indicative of the extent of international capital mobility before and after 1972.

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