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3.4.1989

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SAVING, INVESTMENT AND THE CURRENT ACCOUNT:
A REVIEW OF RECENT EVIDENCE*

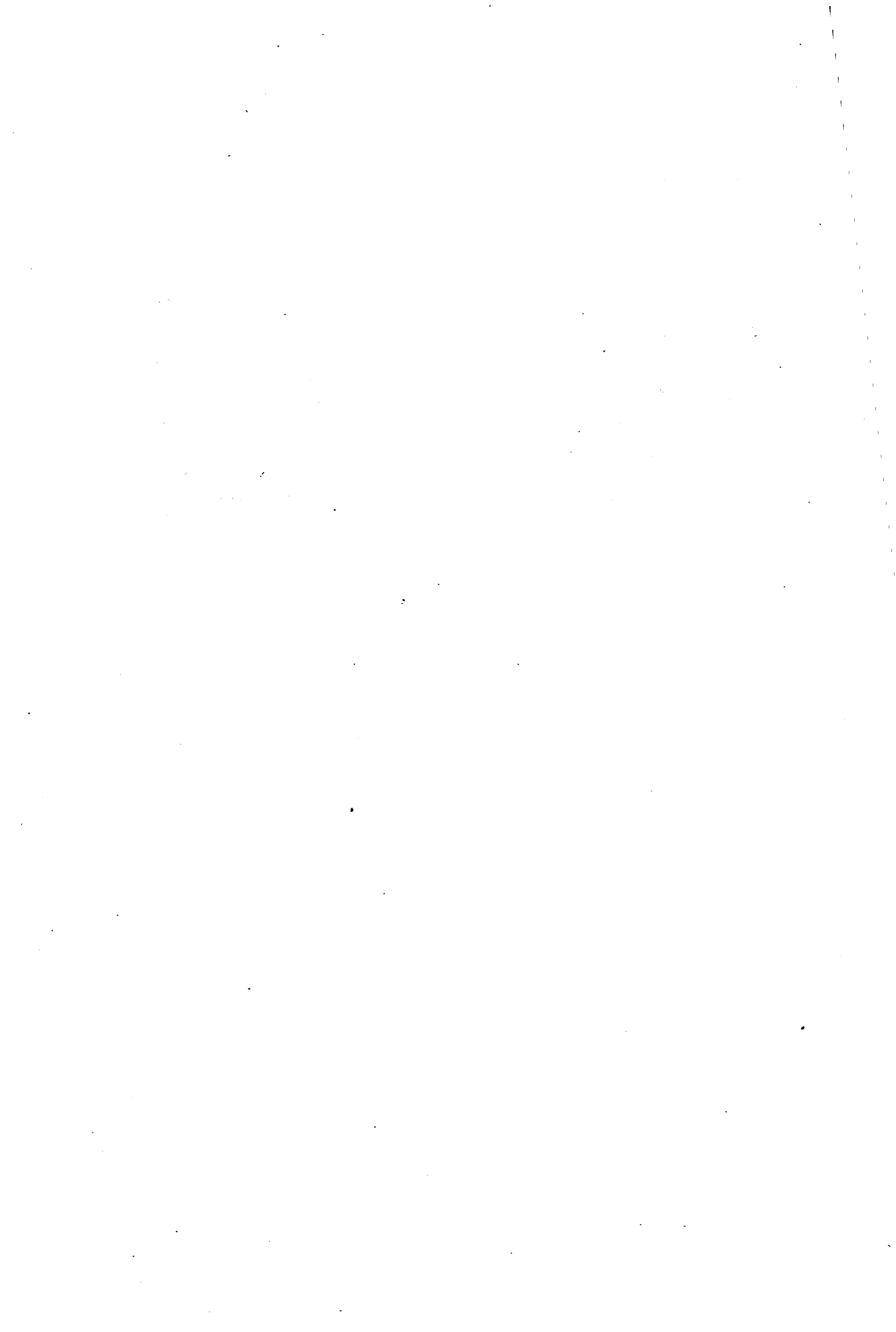
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ABSTRACT

This paper investigates the determination of the current account. In particular, the role of aggregate saving shocks is analyzed using the saving-investment approach as a general frame of reference. Emphasis is paid to testing the Ricardian equivalence proposition. Testing is carried out using cross-section data from 52 countries. Empirical results are clearly at variance with the Ricardian approach. By contrast, the results demonstrate the importance of persistence effects as well as traditional life-cycle model effects.



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1 INTRODUCTION AND THEORETICAL CONSIDERATIONS

This paper deals with cross-country differences in current accounts. More specifically, it scrutinizes the relationship between aggregate saving and aggregate investment. This relationship is utilized both in testing the hypothesis that capital is not perfectly mobile between countries and in deriving a model for cross-country variations in current accounts. This is, in essence, based on the idea, advocated e.g. by Feldstein, that saving shocks are the main determinant of current account variations. As an alternative, we consider a situation where the capital market is more or less perfect and the current account is determined by both saving and investment shocks.

As mentioned above, the subsequent empirical analyses make use of the saving-investment approach which was originally presented by Feldstein and Horioka (1980) and which has since been used on various occasions. This approach suggests the following estimating equation as a basis for testing the role of capital market imperfections:

$$(1) \quad i = a_0 + a_1 s + e,$$

where $i = I/Y$ denotes gross investment, I , in relation to Gross Domestic Product, Y , $s = S/Y$, where S denotes gross saving, and, finally, e denotes the error term. If there are no capital market imperfections (and no policy targets in terms of the current account), i and s should be completely unrelated and, thus, a_1 should be zero. By contrast, if imperfections do exist we should end up by having more or less perfect equality between i and s and, thus, a_0 and a_1 should be zero and one, respectively.

The saving investment approach provides a simple way of analyzing cross-country differences in terms of the current account. An

obvious problem with this approach is the interpretation of causality. Thus, following e.g. Feldstein (1980), one can take it for granted that basically saving constrains investment and thus current account variations are mainly caused by saving shocks. Alternatively, one can assume that saving behavior is very stable and thus that mainly investment shocks cause current account distortions (cf. e.g. Sachs (1981)). It is certainly very difficult to discriminate between these two, rather informal, views, and that is not, in fact, the purpose of this study. Rather, we adopt the Feldstein view and examine whether the basic determinants of aggregate saving help to explain cross-country differences in current accounts.

One particular hypothesis in which we are interested here is the Ricardian equivalence hypothesis strongly advocated by Barro (1974). The Ricardian approach amounts to the statement that the government's fiscal impact is summarized by the present value of its expenditure and thus budget deficits have no first-order effect on the economy. In particular, there is no effect on aggregate saving nor any effect on the current account.

Empirical testing makes use of the standard life-cycle model, which in practice means estimating the following saving {s} and current account {ca} equations (see, e.g., Modigliani (1970) for details):¹

$$(2) \quad s\{ca\} = b_0 + b_1s70 + b_2g + b_3y\$ + b_4dem + b_5def + u_t$$

where ca denotes the current account surplus in relation to Gross Domestic Product (i.e. $ca = s - i$), s70 = the lagged aggregate savings ratio, g = the growth rate of real GDP, y\$ = GDP per capita in U.S. dollars, dem = a (vector of) demographic variable(s), def = the government deficit in relation to GDP and u = the error term. As far as the demographic variables are concerned, three proxies are used: dep = population aged 0 - 14 in relation to total population, ret = population aged 65 and over in relation to total population and mle = expectation of life at birth. In addition to these explanatory variables, a government expenditure variable ps (defined

as public consumption in relation to GDP) was included in the model to account for possible "government dissipation" effects. Because this variable turned out to be completely insignificant in all estimations it was dropped from the final specification.²

2 EMPRICAL RESULTS

Equations (1) and (2) are estimated using cross-section data from 52 countries. The data are annual and correspond to the year 1985. As far as the lagged values are concerned, they are derived from the year 1970. The growth rate of real GDP is computed over the five-year period 1980 - 1985. The main data source is the United Nations' National Accounts Statistics. The demographic data are derived from the United Nations' Demographic Yearbook. The data for the deficit variable are derived either from the United Nations' National Account Statistics (UN) or from International Financial Statistics (IFS). The U.N. concept of deficit corresponds to general government net saving while the IFS concept corresponds (roughly) to general government net lending. Unfortunately, the data for the deficit variable could not be derived for all sample countries. The respective subsamples consist of 35 (UN) and 43 (IFS) observations. The (ratio of the) key variables I and S for 1970 and 1985 is shown in Figure 1.³

Estimation is carried out using both OLS and Huber's M-estimator (see Huber (1981) for details). The latter is used take to into account possible outliers in the data: both (the Jarque-Bera) test for residual normality and an outlier test (see Cook and Weisberg (1982)) suggest that the data do indeed contain such observations, particularly for the 1970 cross-section. Equation (1) is estimated separately both for 1970 and 1985 to allow for parameter and/or regime comparisons.

Turning now to the estimation results reported in Table 1, the results with both OLS and Huber's M-estimator clearly indicate that, first of all, the correlation between s and i is far from perfect. Thus, the parameter restriction $a_0 = 0$, $a_1 = 1$ can be decisively

rejected. Moreover, it turns out that the relationship between i and s has changed somewhat over time so that the coefficient of s is smaller for the 1985 cross-section than for that of 1970. Thus, if aggregate saving really does constrain aggregate investment, its effect has become smaller over time.⁴

As far as the savings ratio equation is concerned, one can immediately notice that the aggregate saving behavior is very persistent: the coefficient of the lagged (aggregate) savings ratio variable clearly deviates from zero. The same is true with respect to the "income growth" variable g . Moreover, the corresponding coefficient is of reasonable magnitude and correct sign. The coefficient of the income level variable y^* - in the same way as the coefficient of the demographic variables - can be estimated only very imprecisely. Only the coefficient of ret , which denotes the portion of elderly people, deviates in some cases from zero (at the 5 per cent level of significance with the t-test). It is also correctly signed, reflecting the fact that an increase in those portions of population engaged in dissaving depresses the aggregate savings ratio (see Modigliani (1970) and Graham (1987) for further evidence). Finally, some comments on the deficit variable merit note. As the results presented in Table 2 indicate, the coefficient of this variable can be estimated reasonably precisely. The sign of the coefficient implies a negative saving effect: a decrease in government saving is not offset by an increase in private saving. This result is clearly at variance with the Ricardian equivalence hypothesis. It is also at variance with some recent empirical findings of Barro (1988).

What about the current account equations? As might be expected, the explanatory power of these equations is much less than that of the savings ratio equations. Moreover, the coefficient estimates suggest that variations in the current account do not only reflect shocks in saving behavior. Thus, the "income growth" variable is now negative probably indicating the positive relationship between investment and growth. Otherwise, the determinants of aggregate saving also affect the current account. Thus, the lagged savings ratio and the income level variable have a positive effect and the portion of elderly

people a negative effect on the current account. Perhaps more importantly, the government deficit has a negative effect on the current account, which suggests that the recent U.S. experience is not an isolated exception among countries.⁵ Still, one caveat must be noted. Some of coefficients are not very precisely estimated and hence some caution is required in interpreting the results. Even so, it should be pointed out that, at least, outliers and heteroscedasticity do not seem to represent serious problems. Our final comment concerns the "income level" variable y . Its positive coefficient estimate suggests that current account deficits are more typical for developing countries. Given the results with the savings ratio equation, one is tempted to interpret this result as evidence for better investment opportunities in the LDC's rather than deficient domestic saving in these countries. Thus, we can conclude that cross-country differences in current accounts can be reasonably well analyzed and explained using the saving-investment approach. Empirical analyses with this approach suggest that the "perfect capital market" assumption is not, after all, at variance with empirical evidence, contrary to what e.g. Feldstein has argued.

Altogether, it seems obvious that the predictions of the Ricardian equivalence hypothesis are not consistent with the data. And not only is aggregate saving adversely affected by budget deficits but also the current account. One is tempted to argue that, instead, our results give some support for the traditional life-cycle model à la Houthakker and Modigliani.⁶ Another result which is perhaps worth mentioning here is the persistence of saving behavior: changes over time in a cross-section of countries appear to be rather small. This result probably reflects the importance of such institutional factors as the tax system, social security, and so on. Clearly, further analysis is required in this respect.⁷

Table 1 Estimation results of equations (1) and (2)

dep. var.	const.	s	R2	SEE	F	year	method
i	.087 (5.51) [5.92]	.648 (8.15) [8.99]	.571	.036	18.44 (0.00)	1985	OLS
i	.085 (5.20)	.654 (7.90)	.571	.036	18.42 (0.00)	1985	ROB
i	.088 (3.91) [3.78]	.737 (7.13) [7.82]	.504	.050	15.00 (0.00)	1970	OLS
i	.071 (4.62)	.785 (11.11)	.504	.050	14.08 (0.00)	1970	ROB

dep. var.	const.	s70	g	y\$	ret	R2	SEE	method
s	.058 (2.95) [2.70]	.650 (6.09) [5.40]	.994 (4.19) [3.90]	.002 (0.75) [0.75]	-.372 (1.83) [1.96]	.627	.040	OLS
s	.055 (2.85)	.620 (5.96)	1.197 (5.18)	-.000 (0.73)	-.197 (1.00)	.616	.041	ROB
ca	-.078 (4.49) [6.02]	.347 (3.70) [4.20]	-.411 (1.98) [1.93]	.002 (1.16) [1.07]	-.220 (1.23) [1.35]	.345	.035	OLS
ca	-.084 (5.10)	.393 (4.44)	-.514 (2.62)	.001 (0.56)	-.295 (0.94)	.337	.036	ROB

t-ratios are in parentheses immediately below the coefficient estimates; below them are White's t-ratios adjusted for heteroscedasticity; F denotes the F-test statistics for the parameter restriction $a_0 = 0$, $a_1 = 1$ (marginal significance levels are in parentheses); ROB denotes Huber's M-estimator, the estimates are computed using the tuning constant 1.345. The number of observations is 52.

Table 2 Testing the budget deficit effects

dep. var.	const.	s70	g	y\$	ret	def	R2	SEE	data
s	.074 (3.93)	.506 (4.71)	1.443 (6.63)	-.001 (0.27)	.265 (0.13)	-.401 (3.20)	.739	.030	IFS
s	.098 (3.78)	.408 (3.27)	.604 (1.94)	.003 (1.19)	-.226 (0.96)	-.365 (2.45)	.548	.037	UN
ca	-.060 (3.44)	.298 (3.04)	-.224 (1.12)	.001 (0.47)	-.013 (0.13)	-.475 (4.14)	.527	.030	IFS
ca	-.086 (3.21)	.334 (2.59)	-.425 (1.32)	.003 (1.20)	-.170 (0.71)	-.155 (1.01)	.391	.038	UN

The symbols are the same as in Table 1 except for the deficit variable def, which is explained in the text. The number of observations with the IFS data is 43 and with the UN data 35.

FOOTNOTES

- 1 Obviously, if capital were completely immobile, and consequently saving equalled investment, the whole capital account equation would just collapse to some random term. This possibility can perhaps be easily ruled out here.
- 2 "Government dissipation" reflects the difference between resource cost and the market value of government-produced goods and services. Thus, it represents some sort of negative income for households which should raise the household savings ratio (see Kormendi (1983) for details). Alternatively, ps could signal the level of social security which, in turn, should have a negative effect on saving, at least if Feldstein's (1974) "social security depresses saving" argument holds.
- 3 The selection of countries was determined by the availability of (reasonably reliable) data for these two sample years. The year 1970 was chosen because it is typically the first year for which the data are compiled according to the current system of national accounts (i.e. the so-called "New SNA"). A list of countries, together with a description and a printout of the data, is available upon request from the author.
- 4 The degree of correlation has, however, increased somewhat, reflecting the decrease in the cross-country variance of i . The fact that the coefficient a_1 is smaller for the 1985 than for the 1970 cross-section has an obvious mirror image indicating that if s is regressed by i , the respective coefficient estimate would be larger (even if not in a one-for-one relationship because that coefficient would simply not equal $1/a_1$). Our main result is clearly at variance with that of Feldstein and Horioka (1981), who could not reject the parameter restriction $a_1 = 1$ (no restriction with respect to a_0 was tested, and the R^2 turned out to be .91). One may suspect that the explanation lies in the data which Feldstein and Horioka use. Their data are 15- or 5-year averages of i and s for 16 OECD countries over the period 1960 - 1974. Clearly, the cross-country variation of i and s is rather small with these data reflecting, among other things, the fact that investment opportunities do not vary greatly over countries. Thus, in the case of 15-year averages the sample standard deviation is .045 for s and .041 for i - the sample means being .250 and .254, respectively. In our case the corresponding numbers for the 1985 cross-section are: $S.D.(s) = .063$, $Mean(s) = .188$, $S.D.(i) = .054$, $Mean(i) = .209$. More important, however, is the use of averaged data. It is obvious that for periods of five to fifteen years current account behavior is probably more affected by policy actions (according to some current account targeting rule) than by any pure investment or saving shocks. At least, it is difficult to distinguish between the effects of policy and the effects of market imperfections. If, instead, short-frequency data are used, one may safely assume that these policy considerations do not distort the results.
- 5 See, also, McKinnon (1980), who argues forcefully that the government deficit is - at least for Japan and the U.S. - the

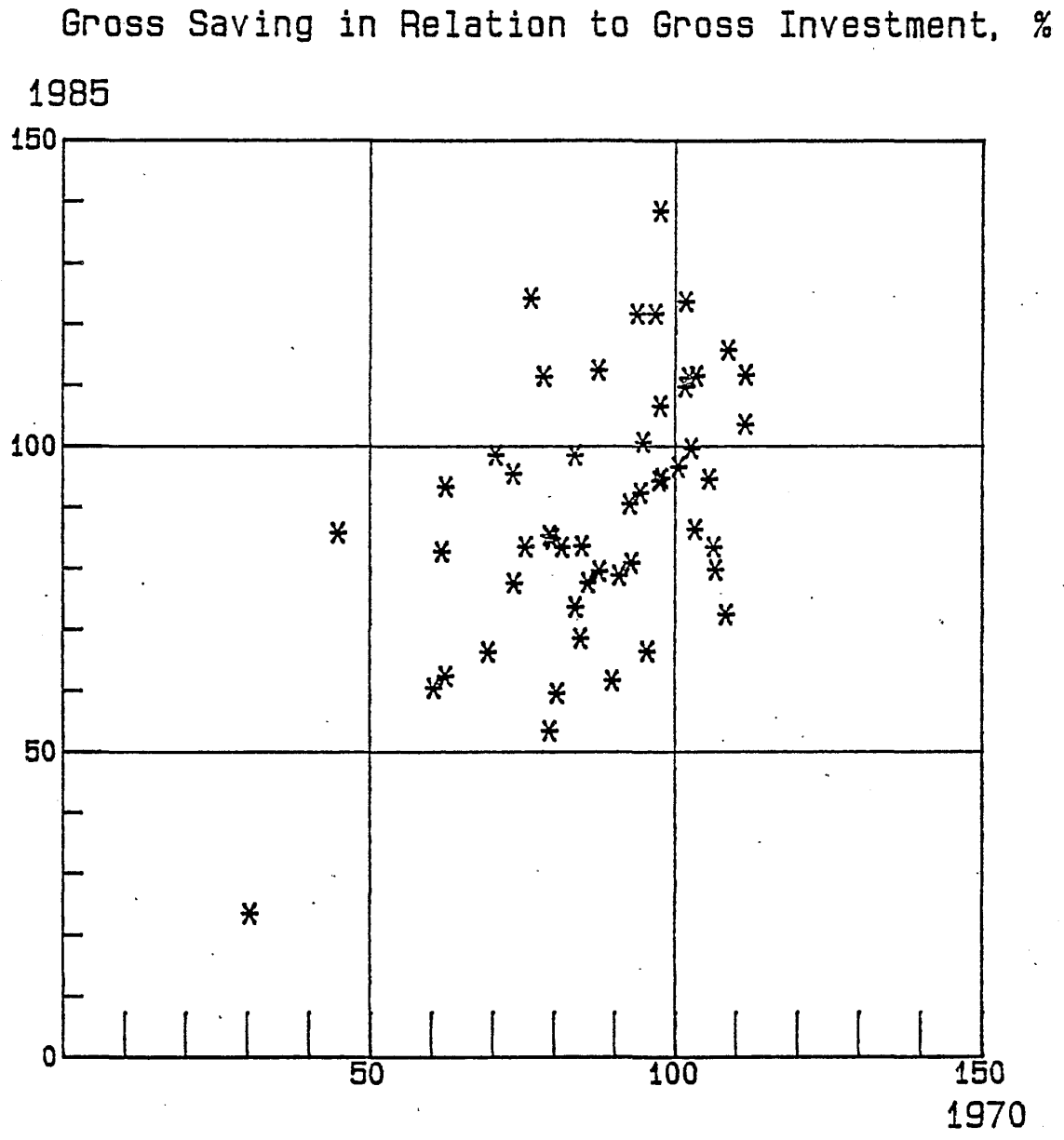
main determinant of the current account. As far as the persistence of deficits is concerned, some idea can be obtained by scrutinizing the scatter plot of the observations presented in Figure 1.

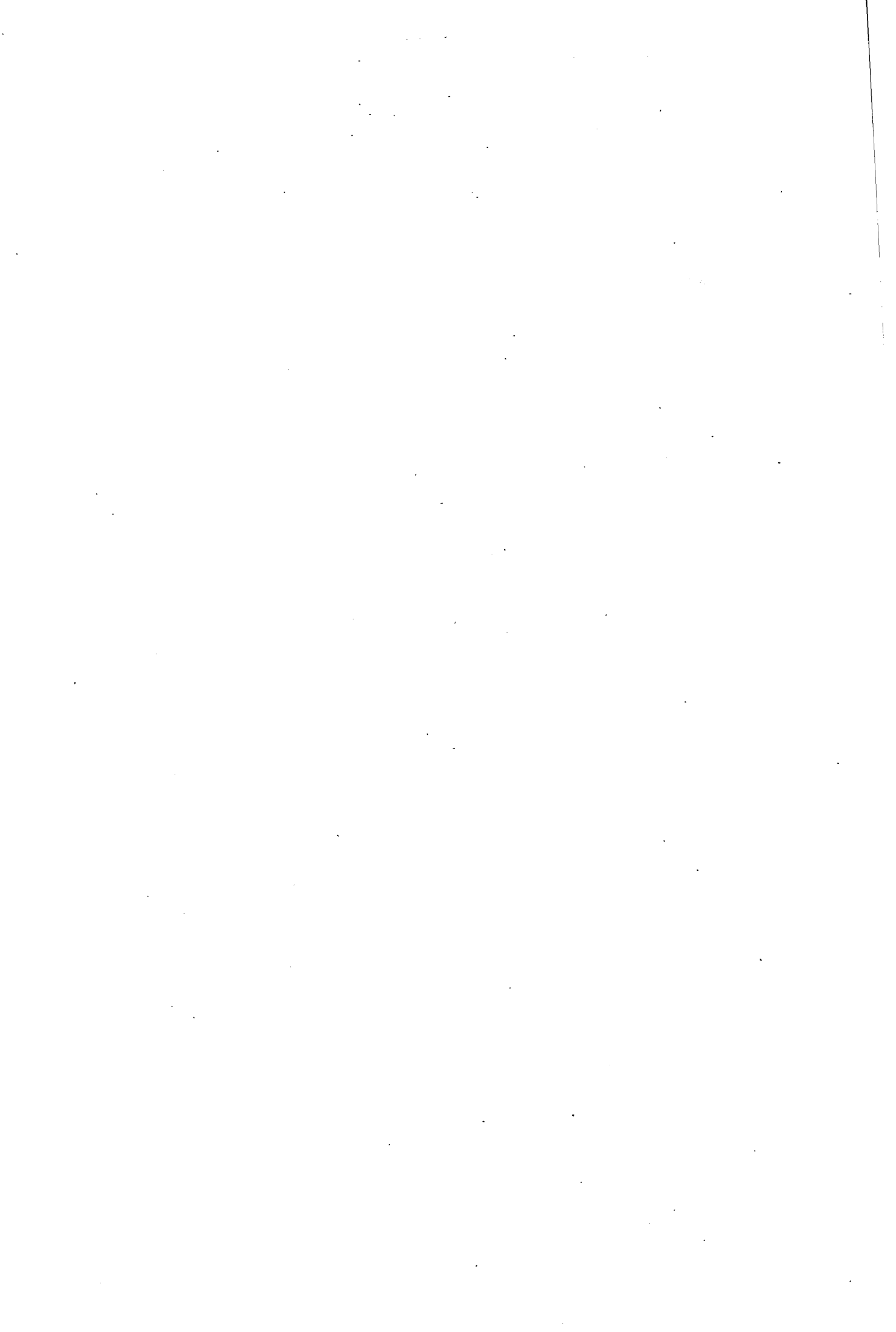
- 6 The question of the performance of the life-cycle model in explaining cross-country differences in savings ratios seems to be somewhat controversial. See, e.g., Graham (1987) and Koskela and Virén (1989).
- 7 See Skjaveland and Stokstad (1988) who use data from 10 OECD countries to demonstrate the importance taxes, social security contributions and the structure of employment.

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FIGURE 1





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