The regime-dependent nature of twin deficits: long-run relation and short-run dynamics across boom and busts

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August 30, 2015

Abstract

We investigate the twin deficits hypothesis using US data. At the low frequency, we find that cointegration evidence in support of the existence of twin deficits weakens during busts and strengthens during booms. Our Markov-switching (MS) analysis shows high frequency regime-dependence of both the twin deficit and the causality direction within it. We identify two well-defined MS regimes: NBER recessions combined with asset price busts, and NBER expansions combined with asset price booms. The global picture provided by our analysis that looks both at long-run and short-run dynamics helps to discriminate among different theories on the twin deficit relationship.

JEL Classification: C32, F4, E62

Keywords: Twin deficits, Cointegration, Markov-switching vector error correction model, Granger causality

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1 Introduction

The empirical evidence on the twin deficit hypothesis - i.e. the positive relation between trade and government deficits - is not easy to detect using standard econometric techniques. One reason is the different behavior of the relationship between the current account balance and the budget balance at short (see Corsetti and Müller, 2006) and long-run frequencies (see Fidrmuc, 2003; Holmes, 2011). These conflicting results could be ascribable to threshold effects behind both the long-run relationship and short-run dynamics or to structural breaks and regime shifts.[1]

We investigate the twin deficits hypothesis employing quarterly US data for the period 1973-2013. Our first task is to differentiate the static long-run relation between trade and government deficits from the dynamic error correction component. Our second objective is to detect non-linearities (through a Markov-Switching Cointegrated VAR (MS-CVAR)) both in the long and short-run relations. In this framework, we want to analyze the links of causality but even how these change between regimes. It is well known that the type of causal link has important policy implications: for example, the possibility for policy-makers to influence the current account through fiscal adjustments depends on the Granger-causality; it is also important on theoretical grounds, because the Granger-causality helps to discriminate between competing theories that are applicable to different countries or different periods.

While recent theoretical and empirical analyses of the relation between the current account and the budget balance suggest that it is subject to structural changes, to our knowledge its regime-dependent nature has not been explored intensively yet.[2]

After a deeper discussion on related literature (Section 2) and some theoretical background (Section 3), we first (Section 4.1) report dynamic trace tests on cointegration between trade and government balances to detect periods where the long-run relation tends to vary. Then, considering the role of other relevant variables we employ first a VECM representation (Section 4.2) and then we detect non-linearities, Granger-causality and its possible change between regimes at high frequency by means of a MS analysis (Section 4.3). Section 5 concludes.

2 Related empirical literature

In the past, a large body of papers studied the relationship between the current account balance (CUR) and the budget balance (GOV) employing a variety of empirical approaches. Some use single equation techniques (e.g. Summers, 1988; Bernheim, 1988; Roubini, 1988; Salvatore, 2006) with not unambiguous conclusions on the quantitative effect of fiscal deficits on trade deficits. Others follow a cointegration approach

[1] Think, for instance, to exchange rate regime changes (Leachman and Francis, 2002) or booms in US investment caused by the "new economy" (Mann, 2002).

[2] An exception is Daly and Siddiki (2009), who explore this issue permitting regime shifts in the long-run relation - but not in the short-run and in the causal links - and employing a single equation approach analysis.

(with or without regime shifts) and/or the Granger causality, sometimes applying the Engle-Granger (1987) two step analysis or applying Johansen's (1988) 'system based' approach to test for cointegration. Daly and Siddiki (2009), for example, analysing 23 OECD countries, find that allowing for regime shifts increases the number of countries where a long-run relationship is detected. Fidrmuc (2003), studying the twin deficit relationship for 18 OECD countries for 1970-2001 with a 1989 break, does not find evidence of cointegration for the second sample in the United States. Holmes (2011) is the first study that examines the threshold cointegration between CUR and GOV within a vector error correction model employing the Hansen and Seo (2002) threshold cointegration approach. He finds some evidence that fiscal balances cause external balances, but only in one of the two regimes. In the other, the error correction coefficient suggests that causality runs the other way. In this respect, Holmes' results are supportive of a bi-directional causality, but it is the threshold that is important when it comes to determining which direction prevails. Within the threshold approach in the twin deficits relationship, Nickel and Vansteenkiste (2008) employ a dynamic panel threshold model for 22 industrialised countries. They find that the relationship turns statistically insignificant when the debt to GDP ratio exceeds 80%. Kouassi et al. (2004) test for Granger non-causality between CUR and GOV. They find evidence of causality (unidirectional or bi-directional) between the twin deficits for some developing countries. Only few studies carry out econometric testing for identifying unknown dates for structural breaks or regime changes: Hatemi and Shukur (2002) identify a structural break in 1989 and find that while before that date causality runs from GOV to CUR, the opposite holds after 1989. Some contributions are also devoted to investigate more specifically causality, outside the original linear Granger causality and cointegration approach of the previously cited literature (see Xie and Chen, 2014). Leachman and Francis's (2002) use the Engle–Granger two-step procedure and Johansen testing and find weak support for cointegration between fiscal and trade deficits after 1974 for the US: their error correction modelling suggests that fiscal balances cause external balances.

A different empirical approach to the problem is founded on the idea that statistical evidence on the twin deficit hypothesis requires identifying fiscal shocks, isolating them from other shocks and testing whether these move the two deficits in the same or opposite direction. Kim and Roubini (2008) follow this strand within a VAR framework, and get a very provocative result: endogenous - mainly GDP driven - movements of GOV and CUR can result in "twin divergence".[3] For instance, the improvement in the fiscal balance between 1992 and 2000, associated to a sharp worsening of the current account, seems to be strictly tied to the activity level: during economic booms output increases and this, in turn, makes the fiscal balance improve (GOV increases thanks to larger revenues) and the current account worsen (CUR decreases if net exports drop). However, even if smaller compared to output shocks effects, they find "twin divergence"

[3] See also Corsetti and Müller (2006).

even considering just "exogenous" fiscal shocks. Our empirical approach will deal with the possible regime-dependence of these endogenous dynamics too.

3 Theoretical background

The co-movements among CUR and GOV rests on the national account identity:

$$Y = C + I + G + X - M \tag{1}$$

where income (Y) - or GDP- is given by the sum of private consumption (C), investment (I), public expenditure (G) and net exports ($X - M$).

Rearranging, including taxes (T):

$$(X - M) = (Y - T - C) - I + (T - G) \tag{2}$$

and dividing every component by Y the four terms above become, respectively:

$$cur = s^P - i + gov \tag{3}$$

where small caps indicate the ratio of the variables to GDP and s^P is the private savings-to-GDP ratio.

Equation (3) suggests that if domestic investment is financed prevalently by private savings (so that $s^P - i \simeq constant$ or $(s^P - i) \sim I(0)$), then cur and gov are "twins" on the basis of the accounting relation.

3.1 Causal relationships between CUR and GOV

The causal link between the two variables can change with regimes, following different patterns widely discussed in the literature:

- GOV Granger-causes CUR. According to the standard Keynesian macroeconomic models (e.g. Mundell–Fleming), an increase in the budget deficit (GOV decreases), either due to lower taxes or to greater public expenditures, increases the domestic interest rate and attracts foreign capitals. This results, under flexible exchange rates, in the domestic country's appreciation and in a worsening of the current account deficit (CUR decreases).[4] This case is known as *the keynesian case* or *the Twin Deficit Hypothesis*.

- CUR Granger-causes GOV. According to the *current account targeting hypothesis*, the government resorts to fiscal policy to eliminate current account imbalances (Summers, 1988).

- No causal relationship. An exogenous decrease in GOV (i.e. a decline in public savings through a tax reduction) will lead to an instantaneous and equal increase in (net) private savings (Barro, 1974) offsetting the effects on CUR. Under Ricardian equivalence, households react to the tax reduction anticipating their future increase

[4] Under fixed exchange rates the fiscal stimulus worsens the current account deficit through an increase in output.

thus raising private savings (rather than increasing net foreign borrowing or current account deficits).

- Bi-directional causality. Twin deficits are associated to the degree of international capital mobility and to the Feldstein-Horioka (1980) puzzle. If domestic saving (both public and private) and investment are not highly correlated (that is, if the Feldstein-Horioka (1980) puzzle does not hold), reflecting high capital mobility, then CUR and GOV tend to move together.

4 Empirical Analysis

4.1 Non stationarity and rolling trace tests

We focus on quarterly data for the US from 1973:1 to 2013:2 to consider the post-Bretton Woods flexible exchange rate period.[5] We first test for the presence of unit roots in *cur* and *gov* and we conclude that both series should be treated as non-stationary and I(1).[6]

As a preliminary analysis we evaluated the rolling trace test on cointegration between *cur* and *gov* considered alone. Figure 1 plots the scaled trace test statistics for the null hypothesis $r = 0$, that is no cointegration, against the alternative of one cointegrating vector, $r = 1$, where r represents the rank of the long-run matrix in the VECM representation.[7] As the Figure shows, rather than cointegrating for the full sample there are periods where the long-run relation tends to strengthen (see e.g. 1978-1980, the second half of the 80s and 90s) and others when it becomes weaker (e.g. 1973-75, the first 80s and 2000-2003). In particular, periods where cointegration weakens tend to coincide with stock market busts[8] and recessions; this tendency reinforces when busts and recessions appear together. Conversely, the cointegration tends to reinforce in periods of expansion and stock market booms.

Looking at equation (3) one can identify the constant in the cointegrating vector with the difference between private savings-to-GDP ratio and private investments-to-GDP ratio ($s^P - i$). Based on that equation we can analyse how net private savings change. Consider the periods when *cur* and *gov* do not cointegrate: these (generally) coincide with the double occurrence of stock market busts and recessions. In this case the government balance should have no long-run link with the current account: move-

[5] The dataset is described in the appendix.

[6] As reported in the first two rows of Table A of the appendix, the unit root null cannot be rejected at the 5% level. Panel b shows that differencing induces stationarity.

[7] The continuous plot of the trace test statistics, for a rolling fixed-length window, provides essential information on the time varying pattern of the cointegrating relation and on its force, expressed by the magnitude of the trace statistics. The test statistics are calculated for a rolling of 36 observations time window (which corresponds to 9 years). The sequences of these statistics are scaled by their 5% critical values. We compute the critical values for the test using p-values by MacKinnon-Haug-Michelis (1999). A value of the scaled test statistic greater then one means that the corresponding null hypothesis can be rejected at the 5% level for the specified sub-sample. Several trials with larger windows and various lags in the VAR specification have been made with similar results.

[8] See Bordo *et al.* (2007) for stock market boom and busts dating.

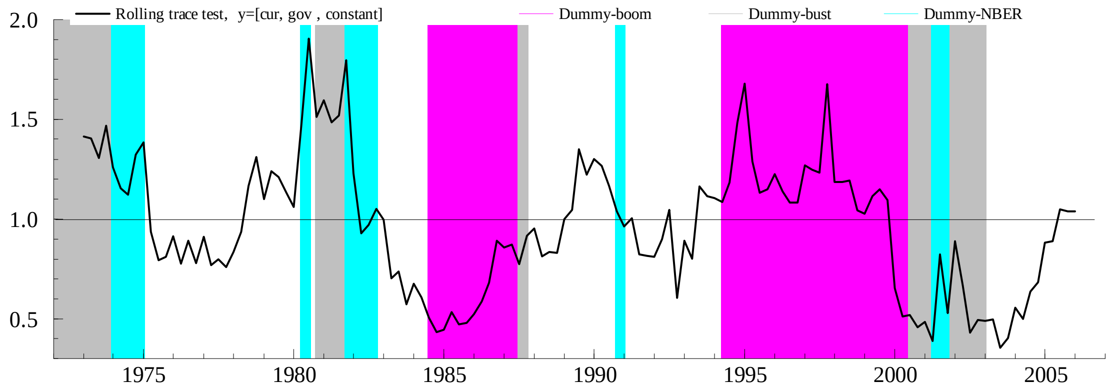

Figure 1: Rolling trace-test (window=9 years) for cointegration between *cur* and *gov*; three dummies, respectively, for financial booms, busts and recessions (Dummy-NBER).

ments in government budget deficit and in net private savings (usually high during recessions) offset each other. This could be the outcome of a Ricardian behavior: take a government who decreases taxes; if ricardian equivalence holds, economic agents, rather than increasing consumption, would decrease it since they expect higher taxes in the future. However, this is even consistent with Summers (1988) who suggested that governments, in order to restrict current account imbalances, may adjust their budget deficit to offset the gap between private savings and investment. A third explanation might come from the traditional crowding out effect of private investments by public expenditures. On the other hand, in periods of expansion, when cointegration strengthens, a change in the government deficit would reflect itself on the current account: in this case the burden of adjustment would not fall on net private savings which are low in expansions, by the way.

To sum up, we find the twin deficit hypothesis to vary with the business cycle: during recessions the cointegration weakens and net private savings tend to increase while the opposite holds during expansions.

4.2 A long-run analysis

Statistical tests presented in the literature (see Fidrmuc, 2003) show that for the majority of the countries, including US, *cur* and *gov* do not show cointegration for long samples; therefore, modelling their relationship requires to consider other variables. In order to test the validity of the twin deficits hypothesis in the context of cointegration theory, empirical studies have typically used a linear model such as

$$cur_t = \beta_0 + \beta_1 gov_t + \beta_2 i_t + u_t \tag{4}$$

where u_t is the error term. Therefore, private propensity to save s^P is usually considered as a stationary variable, not cointegrated with gov and i. Moreover, this equation suggests that in presence of high capital mobility, according to the Feldstein-Horioka puzzle, increasing public savings are expected to mainly affect cur rather than i. $\beta_1 > 0$

and $\beta_2 < 0$ are expected, with an higher magnitude of β_1 (and lower of β_2) in presence of higher capital mobility. Moreover, $\beta_1$depends on the presence and intensity of a "Ricardian effect" of gov on s^P: if changes in public savings are offset by opposite changes in private savings the link between gov and cur (and i,under capital immobility) is breaks down.

Mainly due to the non-stationarity of the private propensity to save, this set of variables turns out to be not appropriate for US data in presence of a large sample. We thus consider a wider set of variables; in particular we include the real rate of interest and real GDP considering a system that is made up of five variables ($cur, gov, i, rgdp$, rir) and whose equilibrium relationship, normalized for cur, is:

$$cur_t = \beta_0 + \beta_1 gov_t + \beta_2 i_t + \beta_3 rgdp_t + \beta_4 rir_t + u_t \tag{5}$$

where $rgdp$ is the log of the real GDP, rir is the 3-month (ex-post) real interest rate. The presence of $rgdp$ is intended to detect the influence of the activity level on the relation between cur and gov and to control for the cyclical components of gov in the short-run analysis, while the real interest rate is a crucial variable of the transmission mechanism of both fiscal, monetary and financial shocks.

Testing for the presence of unit roots in these additional variables we find that these series should be treated as non-stationary and I(1) (see Table A in the appendix).

Given the $I(1)$ nature of the data we assume that the true dynamics can be approximated by a linear VECM(k-1) (Vector Error Correction Model), where the variables are allowed to depart from equilibrium and the adjustment coefficients represent the speed of correction:

$$\Delta y_t = \upsilon + \Pi y_{t-1} + \sum_{j=1}^{k-1} \Gamma_j \Delta y_{t-j} + \varepsilon_t \qquad \varepsilon_t \sim N(0, \Omega) \tag{6}$$

where $y_t = \ (cur_t \ \ gov_t \ \ i_t \ \ rgdp_t \ \ rir_t)'$, υ is the vector of intercept terms, the matrices Γ_j and Π contain, respectively, the short and long-run information of the data, and Ω is the variance-covariance matrix of the errors. The hypothesis of cointegration is formulated in terms of reduced rank restriction on Π, that can be factorized as $\Pi = \alpha\beta'$. Thus, $\beta' y_t$ is the long-run equilibrium relation. In model (6) the variables are allowed to depart from equilibrium and the adjustment coefficients α represent the speed of correction. The traditional linear approach to error correction modelling assumes the time invariance of the speed of adjustment α towards long-run equilibrium and of all the other coefficients $(\upsilon, \beta_i, \Gamma_j, \Omega)$. However, since the evidence on structural breaks and regime shifts in the twin deficit literature suggests that the assumption of parameter constancy is too restrictive, we want to generalize system (6) to account for regime shifts.

Our assumption is that the short-run dynamic behaviour of cur and gov can be characterized by temporary correlations of opposite sign if compared to the long-run correlation, as suggested by the twin divergence hypothesis.

We now estimate and test model system (6). Starting with an unrestricted $VAR(7)$, after tests of model reduction we opt for a $VAR(4)$.[9] We then apply Johansen's (1988, 1991) procedure to test for the reduced rank of $\Pi = \alpha\beta'$, where β is the cointegrating vector. The trace and the maximum eigenvalue tests show one cointegrating relationship, as we reject the null hypothesis of rank zero (see Table 1).

Table 1: Cointegration analysis

rank	Eig.value	Trace	5% Crit.Val.	p-value**	Max-test	5% Crit.Val.	p-value**
$r = 0^*$	0.337	102.496	76.973	0.0002	64.980	34.806	0.0000
$r \leq 1$	0.105	37.516	54.079	0.5975	17.553	28.588	0.6139
$r \leq 2$	0.052	19.964	35.193	0.7290	8.502	22.300	0.9308
$r \leq 3$	0.040	11.461	20.262	0.4983	6.400	15.892	0.7419
$r \leq 4$	0.032	5.061	9.165	0.2768	5.061	9.165	0.2768

Note: * denotes rejection of the hypothesis at the 0.05 level. **MacKinnon-Haug-Michelis (1999) p-values.

Table 2 presents a synthesis of the static equilibrium relation, and short-run dynamics in the linear VECM representation. The estimates suggest a positive relation between cur and gov (with a coefficient of 0.223) providing evidence of a "twin deficit hypothesis" at the low frequency. The LR-test rejects the weak exogeneity hypothesis, except for cur (see Panel b of Table 2).

The other cointegrating coefficients' signs show that both i and $rgdp$ have a negative relation with cur (with coefficients, respectively, of -0.648 and -4.392): a rise in investment demand and in production increases the demand for foreign goods thus worsening cur.[10] On the other hand, the cointegrated coefficient for rir shows a positive relation with cur (with a coefficient of 0.406): a rise in the real interest rate makes net savings increase hence increasing cur. Note, however, that this result is in contrast to the standard keynesian explanation according to which increasing interest rates should attract foreign capital and, by appreciating domestic currency, lead to a current account deficit. The positive relation between the real interest rate and the current account balance could instead signal a specific role for monetary policy. Monetary policy easing, decreasing real rates, makes borrowing cheaper thus spurring demand. In case demand is devoted to foreign goods an increase in the current account deficit occurs (thus cur decreases).

Cointegration analysis is just our first step: starting with the estimated long-run relationship, we will now extend the analysis by including potential regime shifts in the short-run dynamics and in the equilibrium mean.

[9] The lag order is chosen according to the Schwartz criterion (SC) and the absence of autocorrelation.

[10] The proximity to one of the coefficient on i signals that a high proportion of domestic investment is financed from international sources.

Table 2: Multivariate cointegration analysis

Panel a: cointegrated vector β and coefficients of α						
	Const	cur_t	gov_t	i_t	$rgdp_t$	rir_t
β'	-52.72	1	-0.223	0.648	4.392	-0.406
	(5.757)		(0.092)	(0.180)	(0.473)	(0.086)
		$\alpha_{\Delta cur}$	$\alpha_{\Delta gov}$	$\alpha_{\Delta i}$	$\alpha_{\Delta rgdp}$	$\alpha_{\Delta rir}$
α'		-0.033	0.086	0.062	-0.001	0.111
		(0.022)	(0.041)	(0.014)	(0.0004)	(0.040)
Panel b: LR test of restrictions $\alpha_i = 0$: $\chi^2(1)$						
$\alpha_i = 0$		$\alpha_{\Delta cur} = 0$	$\alpha_{\Delta gov} = 0$	$\alpha_{\Delta i} = 0$	$\alpha_{\Delta rgdp} = 0$	$\alpha_{\Delta rir} = 0$
$\chi^2(1)$		2.1968	4.1170	18.710	10.772	7.1538
[p-value]		[0.1383]	[0.0425]*	[0.0000]**	[0.0010]**	[0.0075]**

Note: Standard errors and p-values, respectively, in round and square brackets.

4.3 Short-run dynamics: a Markov-Switching Analysis

In this section we want to investigate whether the twin deficit hypothesis, found at the low frequency (long-run) movements once the short-run dynamics has exhausted its effects, holds even considering short-run dynamics and, in particular, taking account of nonlinearities.

A MS-VECM is proposed that generalizes model 6 to account for nonlinearities. In this framework we use the procedure introduced by Krolzig (1997), which consists of a two-step approach: the first corresponds to a cointegration analysis in a linear model, the second applies the MS methodology.[11] The model describes the stochastic process that determines the regime-switch by means of an ergodic Markov chain defined by the following constant transition probabilities:

$$p_{i|j} = \Pr(s_{t+1} = i \,|s_t = j), \qquad \sum_{i=1}^{M} p_{i|j} = 1, \quad i,j \in \{1,...,M\} \tag{7}$$

where s_t follows an irreducible ergodic M-state Markov process.

The Davies (1987) upperbound LR-test rejects the linear model and our final model allows for regime shifts in $(\upsilon, \alpha, \Omega)$. Table 3 and Figure 2 and 3 report the results of the two-state MS-VECM[12]:

$$\Delta y_t = \upsilon(s_t) + \alpha(s_t)\beta' y_{t-1} + \sum_{j=1}^{k-1} \Gamma_j \Delta y_{t-j} + \varepsilon_t, \qquad \varepsilon_t \sim N(0, \Omega(s_t)) \tag{8}$$

[11] Saikkonen (1992) and Saikkonen and Luukkonen (1997) show that most of the asymptotic results of Johansen (1988, 1991) for estimated cointegration relations remain valid even in the presence of regime-switching.

[12] As a first step we search for the existence of three regimes. Recently, Psaradakis and Spagnolo (2006) have suggested that a joint determination of the lags order and number of regimes can be done through the information criteria such as SC or AIC. Moreover, the simulations in Awirothananon and Cheung (2009) suggest that SC works better for this purpose. In our case, SC excludes the presence of a third regime.

Table 3: MS(2)-VECM(1)

Eq.	$\upsilon(1)$	$\upsilon(2)$	$\alpha(1)$	$\alpha(2)$	Δcur_{t-1}	Δgov_{t-1}	Δi_{t-1}	Δrgdp_{t-1}	Δrir_{t-1}
Δcur_t	1.440	2.300	-0.027	-0.045	-0.019	0.021	-0.019	-15.05	-0.007
p-value	(0.616)	(0.048)	(0.645)	(0.050)	(0.830)	(0.674)	(0.879)	(0.003)	(0.907)
Δgov_t	-8.073	-5.484	0.152	0.111	-0.255	-0.094	0.370	13.391	0.091
p-value	(0.000)	(0.021)	(0.000)	(0.021)	(0.075)	(0.377)	(0.122)	(0.000)	(0.333)
Δi_t	-2.597	-3.438	0.048	0.068	-0.117	0.069	0.027	15.067	-0.069
p-value	(0.047)	(0.001)	(0.061)	(0.001)	(0.023)	(0.019)	(0.790)	(0.000)	(0.041)
$\Delta rgdp_t$	-0.131	0.065	0.003	-0.001	-0.003	0.0001	0.008	0.090	-0.0007
p-value	(0.001)	(0.000)	(0.000)	(0.000)	(0.030)	(0.916)	(0.000)	(0.313)	(0.592)
Δrir_t	-36.684	0.412	0.732	-0.010	-0.130	0.008	0.313	7.156	0.356
p-value	(0.000)	(0.794)	(0.000)	(0.756)	(0.351)	(0.928)	(0.180)	(0.000)	(0.000)

Note: Linearity LR-test $\chi^2(27) = 2568.4[0.0000]^{**}$ approximate upperbound: $[0.0000]^{**}$.
The p-values in parentheses are based on robust standard errors.

where the cointegrating relation is included as exogenous variable. Hansen and Johansen (1998) have shown that shifts in υ can be decomposed into shifts in the mean of the equilibrium and shifts in the short-run drifts of the system. Therefore, in order to account for variation in β_0 in the MS-VECM estimation, the intercept is not restricted to lie in the cointegration space. Matrices α and Γ are the basis for the Granger-causality analysis.[13]

Two interesting and well defined regimes emerge. As Figure 2 and 3 show, Regime 1 captures all the post 1973 NBER recessions dates that also include stock market busts (with the exception of 1987 financial crisis that does not meet a recession) while Regime 2, the temporally prevailing one (with a probability of remaining in that regime equal to 0.936 against 0.843 of remaining in Regime 1, see Panel b in 4), is characterized by expansion periods and includes the dates of stock market booms.[14]

Furthermore, while Regime 1 includes periods of high volatility of all variables, in particular of the real interest rate, the opposite holds for Regime 2 that includes the Great Moderation years indeed (see the variances in Table 4, Panel a).

From the fourth column in Table 3 we get that Regime 1 is strongly characterized by the adjustment of the real interest rate to the equilibrium error (with a coefficient equal to 0.732), while cur (and to a lesser extent i) does not adjust, thereby being weakly exogenous (the coefficient on cur, equal to -0.027, is not significant). Conversely, looking at the fifth column in the same Table, in Regime 2 the sole variable weakly exogenous is rir (with a non-significant coefficient equal to -0.01).

For each regime we want to establish the following facts on cur and gov: a) their possible weak exogeneity; b) their causality nexus; c) their reaction to $rgdp$; d) their reaction to the real interest rate.

[13] According to Engle and Granger (1987).

[14] As identified in Bordo et al. (2007) and Jansen and Tsai (2010).

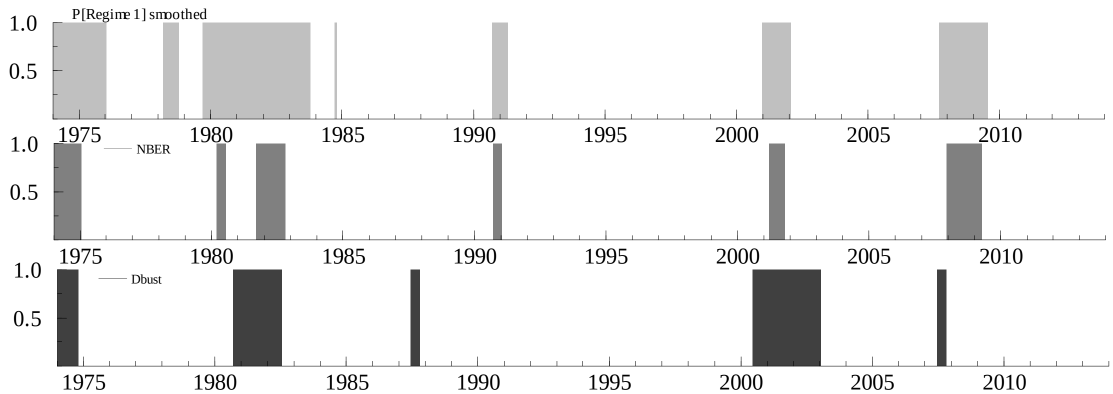

Figure 2: Regimes 1 with NBER recessions and stock market busts.

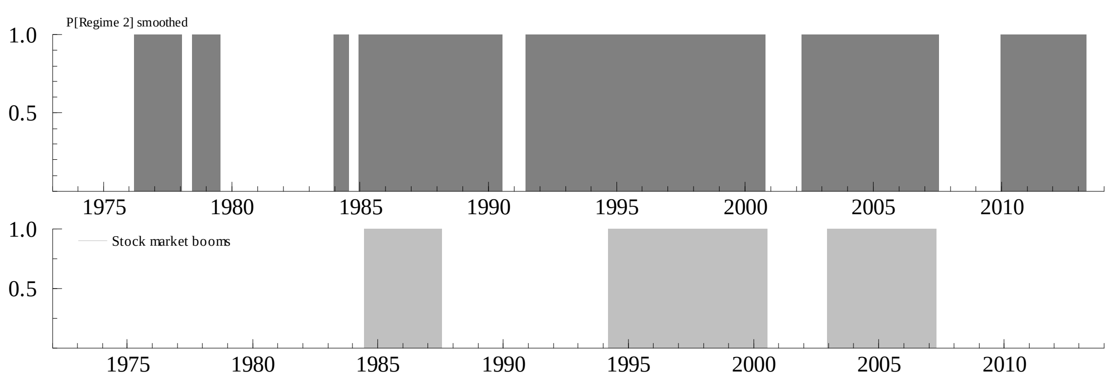

Figure 3: Regimes 2 with stock market booms.

Table 4: Regime Properties

Panel a: Regime-switching variance-covariances matrices					
Regime 1	Δcur_t	Δgov_t	Δi_t	$\Delta rgdp_t$	Δrir_t
Δcur_t	0.215				
Δgov_t	0.0121	0.910			
Δi_t	-0.021	0.141	0.088		
$\Delta rgdp_t$	-0.001	0.005	0.002	0.0001	
Δrir_t	-0.063	0.201	0.158	0.002	1.269
Regime 2	Δcur_t	Δgov_t	Δi_t	$\Delta rgdp_t$	Δrir_t
Δcur_t	0.076				
Δgov_t	0.023	0.221			
Δi_t	-0.007	0.010	0.036		
$\Delta rgdp_t$	0.0001	0.0002	0.0001	0.00002	
Δrir_t	0.006	0.028	0.0004	0.0002	0.092

Panel b: The transition probability matrix		
	Regime $1, t$	Regime $2, t$
Regime $1, t+1$	0.84306	0.06313
Regime 2, $t+1$	0.15694	0.93687

From Table 3 we can infer, for Regime 1, the following features:

a) weak exogeneity of *cur* while *gov* reacts to the equilibrium error;

b) *cur* Granger-causes *gov* (as the adjustment coefficient 0.152 is significant) and this is in accordance with the Current Account Targeting Hypothesis;

c) *cur* reacts negatively only to the growth rate of *rgdp* in the short-run (with coefficient equal to -15.05) but it results not to be reactive to the level of income; *gov* reacts positively both to the *rgdp* level (as the adjustment coefficient 0.152 is significant) and to its growth rate (with coefficient equal to 13.391);

d) *cur* Granger-causes *rir* (as the adjustment coefficient 0.732 is significant) while there is a bi-directional causality nexus between *gov* and *rir* (as both the adjustment coefficients 0.152 and 0.732 are significant).

In Regime 2:

a) both *cur* and *gov* react to the equilibrium error;

b) there is bi-directional causality running from *gov* to *cur*, and vice versa (as both the adjustment coefficients -0.045 and 0.111 are significant);

c) *cur* reacts negatively both to the level of *rgdp* (with coefficient -0.045*4.392) and to its growth rate (with coefficient -15.05); *gov* reacts positively both to the level of *rgdp* (with coefficient 0.111*4.392) and to its growth rate (with coefficient 13.391); therefore, in this regime, there is evidence in favor of a "short-run twin divergence";

d) weak exogeneity of *rir* (as the adjustment coefficient -0.010 is not significant): it Granger-causes both *cur* and *gov* but the reverse is not true.

To sum up, during asset booms and "normal periods" (Regime 2), despite a bi-directional causality between *gov* and *cur*, there is a push for "twin divergence" (see Kim and Roubini, 1988) stemming from GDP that is found to be positively related to *gov* and negatively to *cur*. This works in the process of error correction to the long-run equilibrium, but it is not strong enough to produce a long-run signal. This output effect is asymmetric across the cycle since in Regime 1 *cur* does not react to the *rgdp* level. During busts (Regime 1), both the behaviour of government deficits and of the real interest rate are Granger-caused by concerns on the external balance. The weak exogeneity of *cur* under Regime 1 implies that current account imbalances are not ascribable to the fiscal balance (or to any other variable included in our analysis). Rather, current account imbalances could be the consequence of a "flight to quality" phenomenon: since the U.S. recessions included in our sample coincide with world-wide ones, U.S. dollars, in these periods, are considered as a "safe asset" from overseas investors. The increased demand for this currency, causing an exchange rate appreciation, worsens *cur*. On the contrary, in Regime 2, it is the real interest rate to be weakly exogenous. This result seems to be consistent both with a specific role for monetary policy to explain expansionary phases and, for the more recent years, with Bernanke (2005)'s story according to which a global saving glut has lead to current account deficit in the U.S. as an endogenous reaction to the low real rates that, spurring borrowing, stimulated asset price booms and lowered savings.[15]

5 Conclusion

The picture that emerges from our empirical evidence suggests the existence of a long-run relation between the current account and the fiscal balance that tends to vary with the business cycle: it strengthens during expansions and asset price booms and it weakens during recessions and busts. When these balances are jointly estimated with investment, real GDP and the real interest rate, there is evidence in favour of the "twin deficit hypothesis" at the low frequency (long-run) movements. Therefore, a positive correlation between the two deficits tends to be established in the US experience, once the short-run dynamics has exhausted its effects and only the static equilibrium relation is considered.

Looking at short-run dynamics, employing a Markov-switching analysis, we identify two well-defined MS regimes: NBER recessions combined with asset price busts, and NBER expansions combined with asset price booms. Therefore, as in the long run, we find high frequency regime-dependence of the twin deficit.

The global picture provided by our analysis that looks both at long-run and short-run dynamics can be of help to discriminate among different theories on the twin deficit relationship.

With the double occurrence of stock market busts and recessions we found, in the long-run analysis, that the government balance has no long-run link with the current account. Among the three different explanations put forth for this event (see Section 4.1), that is Ricardian equivalence, current account targeting hypothesis and crowding

[15] For the well-known negative correlation between house price dynamics and current account balances see Bernanke (2010) and Ferrero (2013).

out effect, our Markov-switching analysis points to the second: government concerns for external balances, as suggested by Summers (1988).

On the other hand, in periods of expansion, the long-run analysis finds that cointegration strengthens: a change in the government deficit reflects itself on the current account signalling a limited role in the behavior of this relationship for net private savings, usually low in expansions. The Markov-switching analysis confirms this result during booms since it finds a bi-directional causality between fiscal and current account balances. What is more, it points to the central stage the role of the real interest rate in the explanation because it spurs borrowing, lowers savings and stimulates asset price booms leading to current account imbalances.

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Computations were carried out using MS-VAR codes in Ox (see Doornik, 2013; Krolzig, 1997).

Appendix

Data set and unit root tests

We focus on quarterly data for the United States from 1973.1 to 2013.2. The period was chosen in order to refer the empirical analysis to the post-Bretton Woods period of flexible exchange rates. Figure A presents the relevant dataset in our analysis. Where $cur = (CUR/NGDP) * 100$, $gov = (GOV/NGDP) * 100$, $i = (I/NGDP) * 100$, $rir = R3m - \pi$. CUR = Total Current Account Balance (Source: OECD); GOV =Federal Government Budget Surplus/Deficit; I =Fixed Private Investment; $rgdp$ and $NGDP$ are, respectively, the log of the real GDP and the nominal GDP (All series are seasonally adjusted annual rates. Source: BEA); $R3m$ = 3-Month Treasury Bill rate (Source: Fed's Board of Governors). For the inflation rate π we use the GDP Implicit Price Deflator.

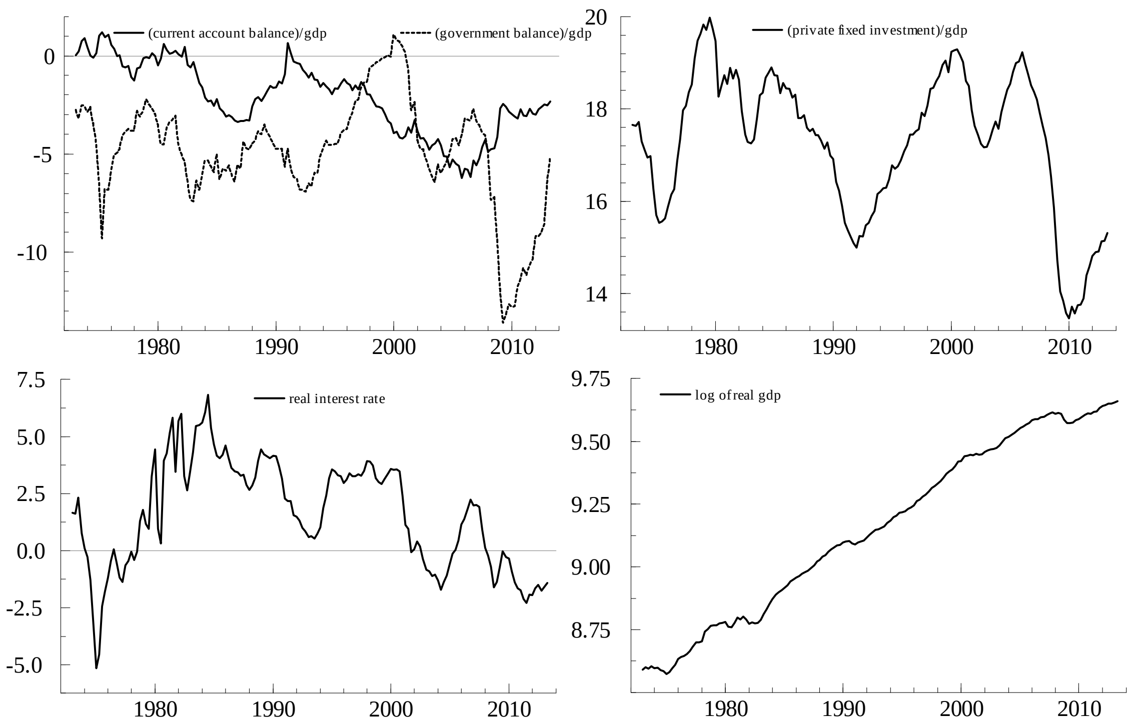

Figure A: Dataset

To check the stationarity of the series we test for the presence of unit roots by the augmented Dickey-Fuller (1979) test (ADF), the Elliott-Rothenberg-Stock (1996) test (DF GLS), the Phillips-Perron (1988) test (PP), and the Kwiatkowski-Phillips-Schmidt-Shin (1992) test (KPSS). All results are reported in Table A, the unit root null cannot be rejected at the 5% level. Moreover, the Kwiatkowski–Phillips–Schmidt–Shin stationarity tests confirm this result.

Table A - Unit root tests

		Panel a		
$Levels$	ADF	$DF-GLS$	PP	$KPSS$
$cv5\%:$	-2.879	-1.942	-2.879	0.463
cur_t	-1.570	-0.899	-1.594	1.787
gov_t	-2.611	-1.727	-2.494	0.564
i_t	-2.776	-1.870	-2.130	0.481
$rgdp_t$	-0.863	2.027	-0.933	3.332
rir_t	-2.348	-1.546	-1.998	0.551
		Panel b		
$Differences$	ADF	$DF-GLS$	PP	$KPSS$
Δcur_t	-12.18	-4.947	-12.18	0.149
Δgov_t	-6.783	-4.342	-11.51	0.056
Δi_t	-4.776	-4.790	-8.223	0.055
$\Delta rgdp_t$	-8.557	-7.642	-8.795	0.194
Δrir_t	-6.895	-7.025	-10.73	0.074

Note: Unit-root tests with intercept term. While ADF, PP, and DF-GLS test the hypothesis of a unit-root, KPSS tests the Null of stationarity against the unit-root hypothesis. The lag order is chosen according to the Schwartz criterion (SC); the PP and the KPSS test are specified using the Bartlett kernel with automatic Newey–West bandwidth selection. Reported critical values: MacKinnon (1996), Elliott-Rothenberg-Stock (1996, Table 1), Kwiatkowski-Phillips-Schmidt-Shin (1992, Table 1).