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IBRAHIM CHOWDHURY
GREGORY GADZINSKI
MATHIAS HOFFMANN

Department of Economics
University of Cologne
Albertus-Magnus-Platz
D-50923 Köln
Germany

http://www.wiso.uni-koeln.de

Asymmetric Dynamics in the Current Account: Evidence from Long-Horizon Data[1]

Ibrahim Chowdhury[a,b,†], Gregory Gadzinski[a] and Mathias Hoffmann[a,b]

[a]University of Cologne

[b]Centre for Financial Research

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Abstract

This letter investigates the presence of asymmetric dynamics in the behaviour of the current account as emphasized in recent theoretical contributions. We estimate a Markov switching model for long-horizon current account to GDP data for six countries and find substantial asymmetries in the behaviour of current account dynamics.

JEL classification: F32; C22

Keywords: Current account dynamics and sustainability, Markov-switching model

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[†]Corresponding author: University of Cologne, Department of Economics, 50931 Cologne, Germany, Email: i.chowdhury@wiso.uni-koeln.de, fax: +49/221/470-5024, tel: +49/221/470-6066.

1 Introduction

Understanding current account dynamics has been the subject of considerable research in the international finance literature. This letter provides a new empirical characterization of the current account behaviour in six industrialized countries by focusing on its asymmetric dynamics, which have lately been suggested in a number of studies (see, for example, Obstfeld and Rogoff, 2000; Taylor, 2002; Chortareas et al. 2004). More precisely, this note is motivated by the idea that a rebalancing of the current account may lead to non-symmetric current account dynamics and differing degrees of sustainability of current account imbalances.

Several explanations are offered to explain asymmetric adjustments in the current account. Recently, transaction costs have been identified as one possible source of asymmetries. For example, Obstfeld and Rogoff (2000) show that even small transaction costs can have large effects on current account dynamics. Taylor (2002) points to the so-called "capital mobility channel" as an additional source of asymmetries, according to which diverse capital mobility policies over time lead to different current account regimes. Moreover, Lane and Milesi-Ferretti (2002) provide some suggestive evidence that country risk and real interest rate differentials are also able to affect current account adjustments across countries. To capture the potential asymmetries we utilize a Markov switching model and analyze current account dynamics for six industrialized countries over the time period 1850 to 2000. We find that the examined countries indeed witnessed substantial asymmetries in their current account dynamics.

2 Modelling Asymmetric Dynamics in the Current Account

An econometric model, which is well able to capture asymmetries in current account dynamics, is a two-state Markov-switching model proposed by Hamilton (1989, 1990), which allows for switching in the mean, the variance, and the autoregressive parameters. More precisely, the Markov switching model takes the following form: $y_t = \kappa'\Gamma$, where y_t denotes the current account-GDP ratio at time t, $\kappa' = [1 - s_t,\, s_t]$ is a vector containing a binary state variable s_t and $\Gamma' = \left[\upsilon_1 + \beta_1' Z + \sigma_1^2 \varepsilon_t,\, \upsilon_2 + \beta_2' Z + \sigma_2^2 \varepsilon_t\right]$; β_i is a $(m \times 1)$ vector of autoregressive parameters β_{ij} for $i = 1, 2$ and $j = 1, \ldots, m$; and Z denotes a $(m \times 1)$ vector of lagged dependent variables; ε_t is assumed to be independently and identically distributed (iid) Gaussian; υ_i and σ_i^2 represent the mean and variance of y_t.

If $s_t = 0$ the time series y_t follows the process described under regime 1: $y_t = \upsilon_1 + \beta_1' Z + \sigma_1^2 \varepsilon_t$. Conversely, under $s_t = 1$, y_t is in regime 2 and characterized by the process $y_t = \upsilon_2 + \beta_2' Z + \sigma_2^2 \varepsilon_t$.

Hamilton (1989, 1990) describes the regime s_t as the outcome of an unobserved discrete-time process, $\Pr(s_t = j \,|s_{t-1} = i) = \rho_{ij}$ for $i, j = 1, 2$, also known as a 2-state Markov chain with transition probabilities $\{\rho_{ij}\}_{i,j=1,2}$.

Since the unobserved regime s_t is presumed to have been generated by some probability distribution, the optimal inference about the current state is based on the history of the observed values of y_t. Thus estimation of the likelihood function is carried out in a recursive fashion by applying the so-called expectations-maximization (EM) algorithm (see, for example, Hamilton, 1990; Krolzig, 1997; Ehrmann, Ellison and Valla, 2003). Each iteration of the EM algorithm consists of two steps. In the expectation step the unobserved states are estimated by their smoothed probabilities for a given set of parameters, here $\theta = \left(\upsilon_i, \beta_i, \sigma_i^2\right)$ for $i = 1, 2$. In the maximization step the parameter set is estimated based on the smoothed probabilities of the last expectation step. This procedure is iterated until convergence is achieved. Hence the EM algorithm provides estimates of the parameter set $\hat{\theta}$ associated with each regime, the transition matrix $\hat{P}$ and the smoothed probabilities.

3 Empirical Results

Annual data on the current account to GDP ratio for Denmark, Germany, Italy, Spain, the UK and the US are obtained from Taylor (2002). For all countries the sample period ends in 2000, while the starting period differs across the countries. More precisely, the sample period starts in 1850 for Spain and the UK, 1860 for Italy, 1869 for the US, 1872 for Germany and 1874 for Denmark.[2] Using long-horizon data has the advantage to rid the analysis from short term nuisance, which may obscure medium- to long-run current account dynamics.

As a preliminary exercise we carried out two types of nonlinearity tests to explore whether the data can indeed be described by a nonlinear process. First, we conducted the RESET test by Ramsey (1969), which tests the null hypothesis of linearity against the alternative hypothesis that the true data generating process is of a general nonlinear nature. The p-values, denoted by ρ_R and reported in Table 1, suggest a clear rejection of the linear null hypothesis. Second, we perform a likelihood ratio test, which specifically tests the null hypothesis of a linear autoregressive process against a Markov switching alternative as specified in the previous section. Notably, for all examined countries the null hypothesis of linearity is strongly rejected against the alternative of a Markov switching model.

Hence, we continue by fitting a Markov-switching specification of the type described above to the current account-GDP ratio. On the basis of conventional lag selection criteria we find a

[2] The statistical reliability of the data has been examined by Taylor (2002).

lag length of one for all countries. The estimated Markov-switching models, reported in Table 1, provide several remarkable insights: All estimated slope parameters are found to be positive, statistically significant, and smaller than unity. The latter result is particularly noteworthy as it implies that the current account to GDP ratio is an overall stationary process (see also Taylor, 2002; Chortareas et al. 2004). We further observe statistically significant switches in the mean and the autoregressive parameters across the two regimes. Except for Italy, we always find a switch in the sign of the mean. In general, the mean in regime 1 takes a negative value, while it is positive in regime 2. For Denmark, Germany, Spain and the UK the slope parameter in regime 2 is larger than in regime 1, while the opposite is true for the remaining countries. The estimated transition matrix $\hat{P}$ provides additional insights on the degree of regime persistence. Both regimes are generally found to be highly persistent, although on average regime 2 appears to be more persistent than regime 1.[3] Except for the US, the probability of a regime switch from state 2 to state 1 is very small and in all cases less than ten percent. Conversely, regime switches from state 1 to state 2 occur at a much higher likelihood, particularly in Germany, Italy and Spain. In the latter country the probability of a regime switch even exceeds the probability of remaining in state 1.

We next shed some light on the identification of each regime. One way to obtain insights on the characteristics of each regime is by examining the plot of the smoothed probabilities of being in one particular state, say state 1, versus the current account-GDP ratio, as displayed in Figure 1. In most of the examined countries, including Denmark, Germany, Italy, Spain and the UK, a significant switch from a current account deficit (surplus) to a surplus (deficit) is frequently accompanied by a fall (rise) in the smoothed probability of being in state 1. These findings may suggest that regime 1 primarily refers to periods, in which the current account is largely in deficit, whereas regime 2 is more closely related to surplus periods. Moreover, the observed switch in the sign of the mean in the estimated Markov switching models also gives rise to this type of regime characterization. Note that in the case of Italy only large current account deficits seem to trigger a switch from regime 2 to regime 1. While this specific identification of regimes may hold for most of the examined countries, it is not entirely applicable to the US. Notably, in the US changes to regime 2 occur less frequently. In fact, a regime switch from state 1 to state 2 is primarily observed for large current account surpluses of about four to five percent.

As a final exercise we compute the half-life to an one percent innovation to the current account-GDP ratio. Except for Italy and the US the half-life recorded under regime 1 is smaller than under the alternative regime (see Table 1). Notably, in Denmark, Germany and Spain the

[3] The elements in the off-diagonal of the transition matrix denote the probabilities of a regime switch, while the elements in the main diagonal reflect the probability that the same state will be maintained.

half-life of a shock in regime 1 is less than one year. Differences in regime-dependent half-lives are most distinct for the UK, where it takes less than two years for a regime 1 shock to be reversed by 50 percent, but almost ten years to achieve a likewise effect for a shock under regime 2. In line with the above mentioned arguments these results would imply that shocks to current account deficits are less persistent than those to current account surpluses. For the US, the half-life in regime 1 is about five years, whereby the half life under regime 2 is somewhat around half a year, suggesting that shocks to large current account surpluses reverse faster, while shocks to current account deficits are longer sustainable. This finding is in line with recent US current account dynamics, which disclose a stronger sustainability on the down-side (see e.g. Obstfeld and Rogoff, 2004).

4 Conclusion

This letter provides empirical evidence that a simple regime-switching model is indeed able to capture asymmetric current account dynamics in several industrialized countries using more than a century of data. Our results are consistent with recent theoretical contributions, which emphasize the presence of substantial asymmetries in the behaviour of current account dynamics and the relative sustainability of current account imbalances in industrialized countries. Overall, the rebalancing of the current account seems to occur faster when the country in question experiences an current account deficit, while the current account adjustment appears more persistent under an initial current account surplus.

References

Chortareas, G.E., G. Kapetanios and M. Uctum, 2004, An Investigation of Current Account Solvency in Latin America Using Non Linear Nonstationarity Tests, Studies in Nonlinear Dynamics and Econometrics, 8(1), 1-17.

Ehrmann, M., M. Ellison and N. Valla, 2003, Regime-dependent impulse response functions in a Markov-switching vector autoregression model, Economics Letters, 78(3), 295-99.

Hamilton, J., 1990, Analysis of Time Series Subject to Changes in Regime, Journal of Econometrics, 45, 35-70.

Hamilton, J., 1989, A New Approach to the Economic Analysis of Non Stationary Time Series and the Business Cycle, Econometrica, 57, 357-84.

Krolzig, H.M., 1997, Markov-Switching Vector Autoregressions. Modelling, Statistical Inference and Application to Business Cycle Analysis, Lecture Notes in Economics and Mathematical Systems No. 454. Berlin: Springer-Verlag.

Lane, P.R. and G. M. Milesi-Ferretti, 2002, Long-Term Capital Movements, NBER Macroeconomics Annual 16, 73-116.

Obstfeld, M. and K. Rogoff, 2004, The Unsustainable US Current Account Position Revisited, Unpublished Paper.

Obstfeld, M. and K. Rogoff, 2000, The Six Major Puzzles in International Macroeconomics: Is There a Common Cause? NBER Macroeconomics Annual 15, 339-390.

Ramsey, J.B., 1969, Test for Specification Errors in Classical Linear Least Squares Regression Analysis, Journal of the Royal Statistical Society, 31, 350-71.

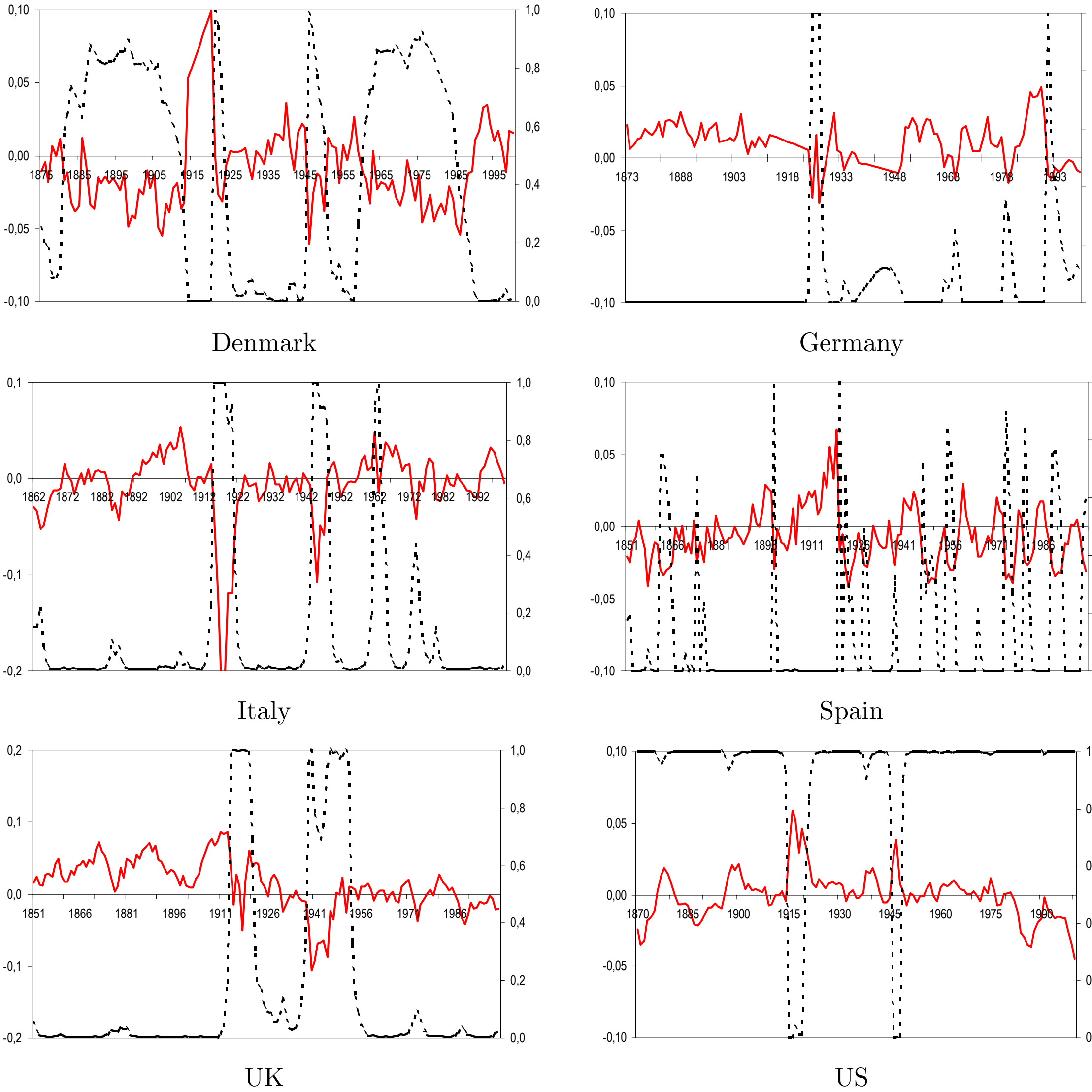
Taylor, A., 2002, A Century of Current Account Dynamics, Journal of International Money and Finance, 21, 725-748.

Table 1: Estimated Markov-Switching Models

	regime 1	regime 2		$\hat{P}$
Denmark	$\hat{y}_t = -\underset{[0.004]}{0.02} + \underset{[0.09]}{0.18} y_{t-1}$	$\hat{y}_t = \underset{[0.002]}{0.004} + \underset{[0.10]}{0.84} y_{t-1}$	$se_1 = 0.026$, $se_2 = 0.001$, $\rho_R = 2.0 \times 10^{-4}$, $\rho_{LR} = 1.0 \times 10^{-5}$, $hl_1 = 0.4$, $hl_2 = 3.9$	$\hat{P} = \begin{bmatrix} 0.88 & 0.12 \\ 0.09 & 0.91 \end{bmatrix}$
Germany	$\hat{y}_t = -\underset{[0.004]}{0.01} + \underset{[0.19]}{0.47} y_{t-1}$	$\hat{y}_t = \underset{[0.001]}{0.003} + \underset{[0.07]}{0.76} y_{t-1}$	$se_1 = 0.011$, $se_2 = 0.001$, $\rho_R = 0.02$, $\rho_{LR} = 1.0 \times 10^{-5}$, $hl_1 = 0.9$, $hl_2 = 2.5$	$\hat{P} = \begin{bmatrix} 0.60 & 0.40 \\ 0.03 & 0.97 \end{bmatrix}$
Italy	$\hat{y}_t = -\underset{[0.02]}{0.02} + \underset{[0.16]}{0.77} y_{t-1}$	$\hat{y}_t = -\underset{[0.001]}{0.001} + \underset{[0.07]}{0.66} y_{t-1}$	$se_1 = 0.045$, $se_2 = 0.001$, $\rho_R = 0.03$, $\rho_{LR} = 1.0 \times 10^{-5}$, $hl_1 = 2.6$, $hl_2 = 1.7$	$\hat{P} = \begin{bmatrix} 0.72 & 0.28 \\ 0.04 & 0.96 \end{bmatrix}$
Spain	$\hat{y}_t = -\underset{[0.001]}{0.03} + \underset{[0.04]}{0.13} y_{t-1}$	$\hat{y}_t = \underset{[0.001]}{0.001} + \underset{[0.06]}{0.73} y_{t-1}$	$se_1 = 0.004$, $se_2 = 0.001$, $\rho_R = 0.01$, $\rho_{LR} = 4.0 \times 10^{-4}$, $hl_1 = 0.3$, $hl_2 = 2.2$	$\hat{P} = \begin{bmatrix} 0.42 & 0.58 \\ 0.09 & 0.91 \end{bmatrix}$
UK	$\hat{y}_t = -\underset{[0.007]}{0.06} + \underset{[0.16]}{0.64} y_{t-1}$	$\hat{y}_t = \underset{[0.001]}{0.001} + \underset{[0.04]}{0.93} y_{t-1}$	$se_1 = 0.034$, $se_2 = 0.012$, $\rho_R = 0.09$, $\rho_{LR} = 1.0 \times 10^{-5}$, $hl_1 = 1.6$, $hl_2 = 9.6$	$\hat{P} = \begin{bmatrix} 0.87 & 0.13 \\ 0.03 & 0.97 \end{bmatrix}$
US	$\hat{y}_t = -\underset{[6.0 \times 10^{-4}]}{0.001} + \underset{[0.04]}{0.87} y_{t-1}$	$\hat{y}_t = \underset{[0.01]}{0.03} + \underset{[0.032]}{0.27} y_{t-1}$	$se_1 = 0.006$, $se_2 = 0.015$, $\rho_R = 0.02$, $\rho_{LR} = 1.0 \times 10^{-5}$, $hl_1 = 5.0$, $hl_2 = 0.5$	$\hat{P} = \begin{bmatrix} 0.98 & 0.02 \\ 0.28 & 0.72 \end{bmatrix}$

Notes: Numbers in square brackets refer to the standard errors of the parameter estimates; se_i denotes the standard error of the regression for regime $i = 1, 2$; p_R refers to the marginal significance level from executing a RESET as described in the text; p_{LR} refers to the marginal significance level from executing a likelihood ratio test; the state-specific half-lives are described by hl_i for $i = 1, 2$, and are constructed as $hl_i = \ln(0.5)/\ln(r_i)$ where r_i denotes the regime-specific root of the Markov switching process. Precise details of the estimation procedure are described in the main text; $\hat{P}$ denotes the estimated transition matrix; entries in the main diagonal of $\hat{P}$ describe the probability that the same state will be maintained, while the off-diagonal elements of $\hat{P}$ describe the probability of a regime switch.

Figure 1: Smoothed Probability versus Current Account-GDP Ratio



Notes: The figure displays plots of the smoothed probability of being in state 1 (dotted line) versus the current account-GDP ratio (solid line).