Monetary Transmission in the Term Structure of Interest Rates in Spain (1995-2003)[1]

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ABSTRACT

The aim of this paper is to investigate whether the effectiveness of the transmission mechanism of monetary policy in Spain has changed since EMU establishment. The analysis is based on the fulfillment of the Expectations Hypothesis under rational expectations and the methodology is implemented through a cointegrated bivariate VAR model. The results reveal the existence of monetary transmission in the term structure in the period prior to EMU, even though the evidence is stronger up to the one-year rate. From 1999, the results are only consistent with a weak evidence of monetary transmission.

Keywords: Interest Rate, Monetary Transmisión, Cointegration, VAR.

Transmisión monetaria en la estructura temporal de tipos de interés en España (1995-2003)

RESUMEN

El objetivo de este trabajo es determinar si han existido diferencias en la efectividad del mecanismo de transmisión de los impulsos de política monetaria en España tras la entrada en vigor de la UEM. El análisis está basado en el cumplimiento de la Teoría de las Expectativas en la estructura temporal de los tipos de interés y la metodología econométrica se implementa mediante un modelo VAR bivariante. Los resultados obtenidos son coherentes con la existencia de transmisión monetaria en el periodo previo a la UEM, siendo menor en los años posteriores a la constitución de la Eurozona.

Palabras clave: Tipos de interés, transmisión monetaria, cointegración, VAR.

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1. INTRODUCCIÓN

The establishment of the Economic and Monetary Union (EMU) in 1999 has meant a drastic change in the economic and monetary conditions of the group of countries that have taken part in the Euro area. The fulfillment of the convergence criteria gave rise to an important range of transformations in the European economies which has not had its equal in recent economic history and, undoubtedly, were it not for the existence and the requirement of the Maastricht Treaty would not have occurred or at least not with the speed they occurred.

The study of the period prior to the setting up of the EMU and the years since it takes on a very remarkable interest for the integrated economies. In particular, and since the EMU is basically a monetary area, it is essential to know the relevance as well as the degree of homogeneity of the monetary transmission process in the different integrating countries.

In view of this, this paper focuses on the Spanish case and tries to analyze whether the effectiveness of the transmission mechanism of the monetary policy in this country has changed after the EMU establishment.

The most recent researches about monetary policy transmission in the EMU[2], carried out using several methodologies, show the relevance of the interest rate channel (IRC) in Europe as opposed to other transmission channels (Angeloni et al., 2003, in a study with data from 1970 to 1999), while in specific countries such as Spain this appears as the dominant channel[3]. Besides, there is evidence supporting that the behavior of European interest rates is consistent with the Expectations Hypothesis (EH) of the term structure at least up to a three-year term (Peersman and Smets, 2001).

In their study with data from 1990 to 2002, Angeloni and Ehrmann (2003) add that monetary transmission has increased through its main channels since 1999, the IRC being responsible for most of these changes as well as for the increase of homogeneity in the European countries. In spite of this, the monetary policy transmission channels do not behave in the same way and with the same intensity in every country.

In this context, it makes sense to go into the peculiarities that the behavior of the term structure of interest rates in Spain is taking on and the changes that have become apparent as a result of the new monetary policy started in 1999.

The methodology used in this paper is based on the contrast of the EH under rational expectations by means of a cointegrated VAR model, in which the spread of interest rates and the changes in the short-term interest rate are the outstanding variables. Empiric evidence in Spain seems to support both the importance of the transmission channel of interest rates (Camarero et al., 2002; Massot and Nave,

[2] See European Central Bank, working papers from 91 to 114, or the compiler book of Angeloni *et al.* (2003).

[3] The IRC is dominant as well in Holland, Portugal, Ireland, Finland and Luxemburg. The dominance is rejected in Germany, France, Belgium, Italy, France, Austria and Greece.

2003) and the existence of a joint dynamic of short-term interest rates up to the one-year term (Goerlich et al., 1995; Barreira et al., 1998, Prats and Beyaert, 1998; Beyaert et al., 2001). However, consensus does not seem to exist about whether the EH is the best theory to explain the yield curve[4].

In line with this research, this paper attempts to obtain suitable evidence of the EH using weekly data for an extensive and complete range of interest rates. It includes interbank money rates from one week to one year and public debt market rates up to ten years. The whole of the data extends from January 1995 to April 2003, and we have distinguished between the period before and after the establishment of the EMU for the purpose of identifying the repercussions that this change in the monetary regime has had in the Spanish term structure of interest rates.

The remainder of this paper is organized as follows. Section 1 summarizes the methodology stated by Campbell and Shiller (1987) for testing the EH under rational expectations. Section 2 describes the data used in the empirical analysis. Section 3 deals with the analysis of cointegration and the choice of the VAR models for the different interest rate spreads. Section 4 presents and discusses the results of the tests that let us to ascertain the validity of the EH in every period and maturity. Finally, the main conclusions are stated in Section 4.

2. ECONOMETRIC METHODOLOGY

Campbell and Shiller (1987) evaluate the EH under rational expectations using a present value model for two variables: the long-term yield, R_t, and the short-term rate, r_t. The model states that R_t, is a linear function of the present discounted value of expected future r_t:

$$R_{t,\infty} = (1-\gamma)\sum_{i=0}^{\infty} \gamma^i E_t r_{t+i} + k \qquad [1]$$

where $R_{t,\infty}$ is the actual interest rate of an asset with a very long maturity; r_t is the short-term interest rate, present and future, maturing in (t+1); E_t represents the rational expectation of the agents conditioned on the information available at time t; γ is a constant discount factor, $0 < \gamma < 1$, defined as $\gamma = 1/1+R^*$, with R^* being an average of the long-run interest rate; and k is a constant liquidity premium.

Equation [1] expresses the basic relationship we are interested in evaluating using data on the Spanish term structure of interest rates.

[4] About the importance of liquidity premium in Spain, see Rico Belda (1999), Pérez et al. (1997), and Flores de Frutos (1995). Against de EH in the debt market in Spain see De Andrés (2004).

The term structure of interest rates could also be formulated in terms of the spread, S_t, between long-term and short-term interest rates:

$$S_t = R_t - r_t \qquad [2]$$

The implications of using the spread in our model [1], instead of R_t, may be illustrated in the following way. If we subtract r_t from both sides of [1] and re-arrange we obtain:

$$S_t = E_t S^* + k \qquad [3]$$

where:

$$S_t^* = \sum_{i=1}^{\infty} \gamma^i \Delta r_{t+i} \qquad [4]$$

The above shows that the spread is the optimal forecast of S^*_t, which in turn is a weighted average of future changes in r_t.

Campbell and Shiller (1987) develop a test of the present value relation that requires cointegration. In particular, if it is demonstrated that the interest rates in levels are I(1) variables, two important consequences are deduced. First, the spread as expresses equation [3] is a linear combination of I(1) variables. Second, according to [4] the spread is a sum of I(0) variables and therefore is I(0). So, if the EH is valid, the long run and short run interest rates combine linearly into a I(0) variable and are therefore cointegrated. This implies, by the Engle-Granger representation theorem, that there exists an Error Correction Model, from which Campbell and Shiller (1987) propose an alternative bivariate VAR model defined on S_t and Δr_t. When the variables are expressed in deviations from their means, this VAR takes the form:

$$\begin{bmatrix} \Delta r_t \\ S_t \end{bmatrix} = \begin{bmatrix} a(L) & b(L) \\ c(L) & d(L) \end{bmatrix} \begin{bmatrix} \Delta r_{t\text{-}1} \\ S_{t\text{-}1} \end{bmatrix} + \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix} \qquad [5]$$

where $a(L)$, $b(L)$, $c(L)$ y $d(L)$ are lag polynomials of order p.

Equation [5] can be written more compactly as:

$$q_t = Aq_{t-1} + w_t \qquad [6]$$

where A is called the companion matrix of the VAR and, for all i, $E_t(q_{t+i}\,|H_t) = A^i q_t$, where H_t is the limited information set containing current and lagged values of R_t and r_t or, equivalently, of q_t.

The implications of the present value relation for the VAR system are the following. A weak implication is that S_t must linearly Granger-cause Δr_t if the spread is an optimal forecast of the future variations of the short-term interest rates. This is a necessary condition that has to be tested on the coefficients of the first equation of the VAR. Also, the sign of the relation between the variables S_t and Δr_t must be positive.

The restrictions of the present value model are, however, more demanding. Projecting equation [3] onto the information set H_t yields the following $2p$ constraints on the coefficients of the VAR:

$$\begin{aligned} a_i &= -c_i,\ i = 1,\dots,p \\ d_1 + b_1 &= 1/\gamma \\ d_i &= -b_i,\ i = 2,\dots,p \end{aligned} \qquad [7]$$

These constraints provide a way of testing the EH using a linear Wald test. With this aim, we can define the auxiliary variable:

$$SR_t = S_t - (1/\gamma)S_{t-1} + \Delta r_t \qquad [8]$$

and thus the Wald test can be carried out over the following regression:

$$\begin{aligned} SR_t = {} & (a_1 + c_1)\Delta r_{t-1} + \cdots + (a_p + c_p)\Delta r_{t-p} + \\ & +(b_2 + d_2)S_{t-2} + \cdots + (b_p + d_p)S_{t-p} + (u_{1t} + u_{2t}) \end{aligned} \qquad [9]$$

3. DATA DESCRIPTION

The data used in this study includes interbank money market rates and zero-coupon bond rates from the Spanish government bond market. The interbank money market rates correspond to Madrid Interbank Offered Rate (MIBOR) from January 1995 to December 1998 and to EURIBOR rates for the period January 1999 to April 2003. Six different maturities have been considered for the short end of the term structure: one week (d07), one month (m01), two months (m02), three months (m03), six months (m06) and one year (y01). The long-term interest rates are zero-coupon rates extracted from coupon-bearing Spanish government bonds[5] using Svensson method and correspond to the following maturities: three years (y03), five years (y05), and ten years (y10).

[5] The database for the estimation of the term structure consists of the daily averages of actual prices crossed by the account-holders in outright spot operations in Spanish government bonds.

From the initial series of daily observations for these rates, weekly series have been obtained by selecting data for every Wednesday, or the nearest trading day if unavailable. The series consist of 52 weekly observations for each year, the result of which is 435 observations per series. All rates are expressed as continuously compounded interest rates for homogeneity purposes. In our opinion, both the quality and frequency of the data are indispensable to obtain consistent results. The short-term rate, r_t, is, in all cases, the one-week rate, and the other eight rates play the role of the long-term rate, $R_{t,i,}$ where, i = m01, m02, m03, m06, y01, y03, y05 and y10.[6] The series of spreads are obtained from [2].

As a result of this data and these definitions, eight models can be defined, whose variable combinations are:

$$(\Delta r_{d07}, \Delta S_{m01}), (\Delta r_{d07}, \Delta S_{m02}), (\Delta r_{d07}, \Delta S_{m03}), (\Delta r_{d07}, \Delta S_{m06}), (\Delta r_{d07}, \Delta S_{y01}),$$
$$(\Delta r_{d07}, \Delta S_{y03}), (\Delta r_{d07}, \Delta S_{y05}) \text{ and } (\Delta r_{d07}, \Delta S_{y10}).$$

4. COINTEGRATION RESULTS AND VAR MODELS

The starting point of the methodology is the analysis of the presence of a unit root in the series R_t and r_t and the stationary of S_t. This is fundamental, as the strategy presented by the Campbell and Shiller (1987) is only valid when the variables R_t and r_t are stationary in first differences and cointegrate, making the spread S_t a stationary variable. The relevant series, expressed in deviations from their means, are displayed in Figures 1 to 3. In particular, Figure 1 shows the interest rates in levels, Figure 2 shows the interest rates in first differences and Figure 3 presents the spreads.

For the unit root analysis we use both Augmented Dickey-Fuller (ADF) test, which detects serial autocorrelation, and Phillips-Perron (1988) test, because of its robustness in the case of heterogeneously distributed errors. In both tests, and for the whole series, we choose the best specification of the auxiliary function, whether including constant and/or trend, or none of them.

The results of the two tests are very similar. This is the reason that in Table 1 we only present the results of Phillips-Perron test because of its better properties. According to Schwert (1989), the test is applied for different values of the truncation parameter. In particular, we choose[7] $q = q_1$, 5, 11, 17, 54, where the first truncation parameter q_1 is defined as q=Ent$[4(T/100)^{2/9}]$, where Ent [z] is the integer part of z and T is the sample size (in our case, T= 435 for the overall period, 209 for pre-EMU period and 226 for post-EMU period). As seen in Table 1, all the interest rates in levels exhibit one unit root for all different q and hence, the interest rates in first differences are stationary. The spread series are also stationary at a significance level of 5% except for the one-week and the one-month series, which reject the null hypothesis of one unit root at 5% but not at 1%.

[6] These terms are equivalent to 4, 8, 13, 26, 52, 156, 260 and 520 weeks, respectively.
[7] See Dolado and Jenkinson (1987), pp.42-43.

According to these results, the short-term and long-term interest rates are cointegrated, as predicted by the theory, and therefore there is a long-term relationship between long and short interest rates (Engle and Granger, 1987). This result could be interpreted as a first evidence of the EH and, therefore, of the existence of monetary transmission in the different cases considered. Finally, the methodology outlined in Section 1 can be implemented.

(a) The eight models are estimated considering from 1 to 52 lags (p) according to VAR methodology and also using OLS over the individual equations of the VAR.
(b) Following, the residuals of these models are analyzed in order to verify the absence of autocorrelation through the Ljung-Box and Breusch-Godfrey (LM) tests considering up to 26 lags (six months), and heteroscedasticity using White test, LM test of ARCH effects (up to order 4, one month) and the existence of serial autocorrelation in the squared residuals (up to 26 lags).
(c) The estimation process has revealed, using both methods, an important degree of autocorrelation and heteroscedasticity in the interest rate series, with significant ARCH effects in all the cases.
(d) In according to these latest results, the equations of the VAR models have been re-estimated using maximum likelihood and introducing GARCH (g, g) effects[8], with g = 1, 2, 3 and 4. The autocorrelation and heteroskedasticity tests of the residuals of the GARCH models have revealed the existence of specifications for each one of the eight pairs of variables in which those effects are not present, and therefore could be chosen.

Once models have been estimated, the selection of the GARCH (g) order and the number of lags (p) for each pair of variables follows these criteria:

(a) First, determine the minimum value for p among the models whose residuals pass autocorrelation tests at 5%.
(b) Second, verify that ARCH effects have disappeared at a level of significance of 5% with the values of p and g selected.
(c) Finally, for the models which accomplish the above criteria, select the specific value for p using the AIC criterion.

As a result, we select one specification for each case and for each period. Table 2 shows the specific values for p, the results of the tests, and the AIC values.

In general, the results show that the GARCH (1,1) specification is the most common in all the cases and periods. However, in pre-EMU period we select a GARCH (2,2) specification for (Δr_{d07} , ΔS_{m01}) and (Δr_{d07} , ΔS_{m06}), and in the post-EMU period we select a GARCH (3,3) specification in (Δr_{d07} , ΔS_{m02}) and (Δr_{d07} , ΔS_{m03}). For the 1995-2003 period, the results are, in general, an average of the results of the pre- and post-EMU periods.

[8] Following Engle (2000), the ARCH correction effect has been done equation by equation.

Finally, in general the number of lags p is higher and more volatile in the second period (with p = 9,11,12,17,26,41) than in the first period (where p = 5,6,9,10,13), which might be explained by the memory necessary to anticipate the changes in interest rates in both regimes.

The existence of this memory, and also the high frequency of the data, explains the large number of lags. This result, together with the low number of observations in our samples, hampers the usage of other econometrical methodologies in this study. Among these methodologies, switching regression models suits our purposes, since they are motivated by the realization that model parameters might change over time according to a set of state variables. These models have an ample history in econometrics (see Maddala and Kim, 1998, for a review) and expand through several directions of interest nowadays. For example, Wong and Li (2001) and Lanne and Saikkonen (2003) have proposed mixtures of autoregressive models with GARCH disturbances that allow regime switches according to a discrete state variable. It might be expected a number of changes in monetary policy during our period of study, linked to the changes in the instruments and application of monetary policy and the functioning of the interbank money market in the first years of EMU. Hence, the extensions of these models through the inclusion of several state variables that influence groups of parameters following the work of Preminger et al. (2007), might limit the growth in the number of parameters that restricts the implementation of switching regression models.

5. TESTING THE IMPLICATIONS OF THE EXPECTATIONS HYPOTHESIS

As stated in Section 1, first and weak evidence of the validity of the EH is that the spread is an optimal predictor of future changes in the short-term rate. That is, S_t should Granger-cause Δr_t in the VAR, and the sign of the relation between the variables S_t and Δr_t must be positive. Further, a Wald test on the coefficients of the VAR provides stronger evidence in favor of the EH.

Table 3 shows the results of these test for all the cases and periods. The Table clearly reveals that the spread Granger-causes short-rate changes, as they should do under the EH, and also the sign of the relation is as expected. These findings are in addition to the existence of cointegration to provide evidence of monetary transmission along the term structure of interest rates both in the pre- and post-EMU periods.

Finally, regarding the Wald tests, it can be observed that the results are quite different in each period. For the pre-EMU period, Table 3 shows the fulfillment of the strong implications of the EH on the coefficients of the VAR models for all the shorter rates from two months up to one year, although not for longer rates. This last result is somewhat paradoxical since, when previous analysis aims at the validity of the EH, there exists an important rejection of the Wald test in all the cases from three to ten years. This situation turns out to be very common in the empirical literature and the most widespread conclusion among the authors[9] is to admit the fulfillment of the EH despite the fact that the different tests point out contradictory conclusions, basically drifted from the Wald test.

Therefore, we conclude that in Spain from 1995 to 1998 there exists monetary transmission in the term structure even though the evidence is stronger up to the one-year rate.

This finding is in harmony with the results contributed by other researchers in which, using various methodologies and periods, the improvement of the IRC in Spain in the years before the EMU is pointed out. For example, for the interbank market, Beyaert et al. (2001) confirm the evidence of transmission in all the interbank interest rates during the period 1989-1997 using weekly data. Barreira et al. (1998) obtain evidence of cointegration and causality in the same terms and in the same market using monthly data. At the long end of the term structure, the study of Camarero et al. (2002) by means of a structural VAR model with monthly data, shows a very significant monetary transmission from 1989 to 1998 using the one-day interbank rate and the ten-year debt rate. Recently, Massot and Nave (2003) obtain results in favor of the EH, and therefore of monetary transmission, in the Spanish public debt market using monthly estimates of zero-coupon rates from Vasicek and Fong (1982) methodology during the period 1993 to 1998.

The success of monetary transmission at this stage was undoubtedly determined by the high integration and expectations of convergence that took place since the entry of the peseta in the European Monetary System[10]. The targets of exchange rates and of stability of interest rates in the financial markets acted as basic elements for monetary transmission.

The results for the post-EMU period reveal significant differences with respect to the pre-EMU period. As shown in Table 3, the Wald tests clearly rejects the strong implications of the EH in any segment of the term structure except for the one year rate, i.e. for the pair (Δr_{d07} , ΔS_{y01}). Although the existence of cointegration and causality among the variables involved in the model led us to affirm the existence of monetary transmission in all the term structure, it is also apparent that the evidence is much weaker than in the years before the EMU.

This result points to a lower effectiveness of the monetary policy of the European Central Bank when compared to the last years of national monetary policy of the Bank of Spain.

Other studies that analyze and compare the effects of transmission in the term structure in individual countries have been scarce until now and when the data of the overall European area is considered, the analysis usually points out evidence of transmission up to the middle segment of term structure. Thus, Angeloni and Ehrmann (2003) obtain clear and important evidence up to the three-year term in Europe. De Bondt (2002) suggests in his study that the movements between the one-

[9] For example, see Shea (1992) and Cuthbertson et al. (1996).

[10] See González (2004) for details about catching up and long-term convergence in interest rates in Spain in the nineties.

day interest rate and other market rates in the Euro area have been getting closer from January 1999, but it is not the case of the long end of the term structure.

The fact that our results contradict the general pattern revealed in these studies points out to the need of analyzing the specific contribution of the EMU to the effectiveness of the monetary transmission mechanism in each of the integrated countries. The favourable valuation of the single monetary policy in the overall area might hide different national appraisals that could be partially related to the divergences in inflation rates[11].

6. CONCLUSIONS

In this paper, the differences in the effectiveness of the transmission mechanism of monetary policy impulses in Spain in the period before and after the establishment of the EMU have been analyzed. The methodology used is based on the contrast of the Expectation Hypothesis of the term structure under rational expectations by means of a cointegrated VAR model in which the spread of interest rates and the changes in the short-term interest rate are the relevant variables. The weekly data used in the analysis have been selected from daily data of interbank rates with different maturities and zero coupon bond rates of the Spanish debt market from January 1995 to April 2003, for the purpose of obtaining evidence for a wide spectrum of maturities.

For both periods and for the whole term structure of interest rates we have found evidence of cointegration and Granger causality, which points to the existence of a certain monetary transmission in Spain. This finding is an important result because the number of studies that show evidence of monetary transmission in both the short and long terms is few. In general, previous works for Spain and Europe support the transmission hypothesis in the short end of the term structure, unlike the empirical work with U.S. data in which it is the long-term which usually admits the fulfillment of the Expectations Hypothesis.

However, the analysis has revealed significant differences in the strength of the interest rate channel in Spain before and after the EMU establishment. In particular, in the pre-EMU period we have found strong evidence in favor of the Expectation Hypothesis in the short-end of the term structure up to the one-year term, which disclose the relevance of the monetary transmission in this segment of the curve. In contrast, in the post-EMU period, the results are according to a weaker monetary transmission in which only the one-year EURIBOR has full valuable informative content about future short-term rates. This result is somewhat surprising in the view of recent studies that show an increase of the interest rate channel power, as well as the rise in the effectiveness and power of other alternative channels of monetary transmission in Europe as a whole. This apparent contradiction should give rise to a starting point for new research underlying the heterogeneity in the financial conditions in which the single monetary policy of the European Central Bank applies.

[11] See, for example, Peersmann (2004), Angeloni and Ehrmann (2004), Ciccarelli and Rebucci (2005), Berben et al. (2005)

FIGURE 1
Interest rates in levels.

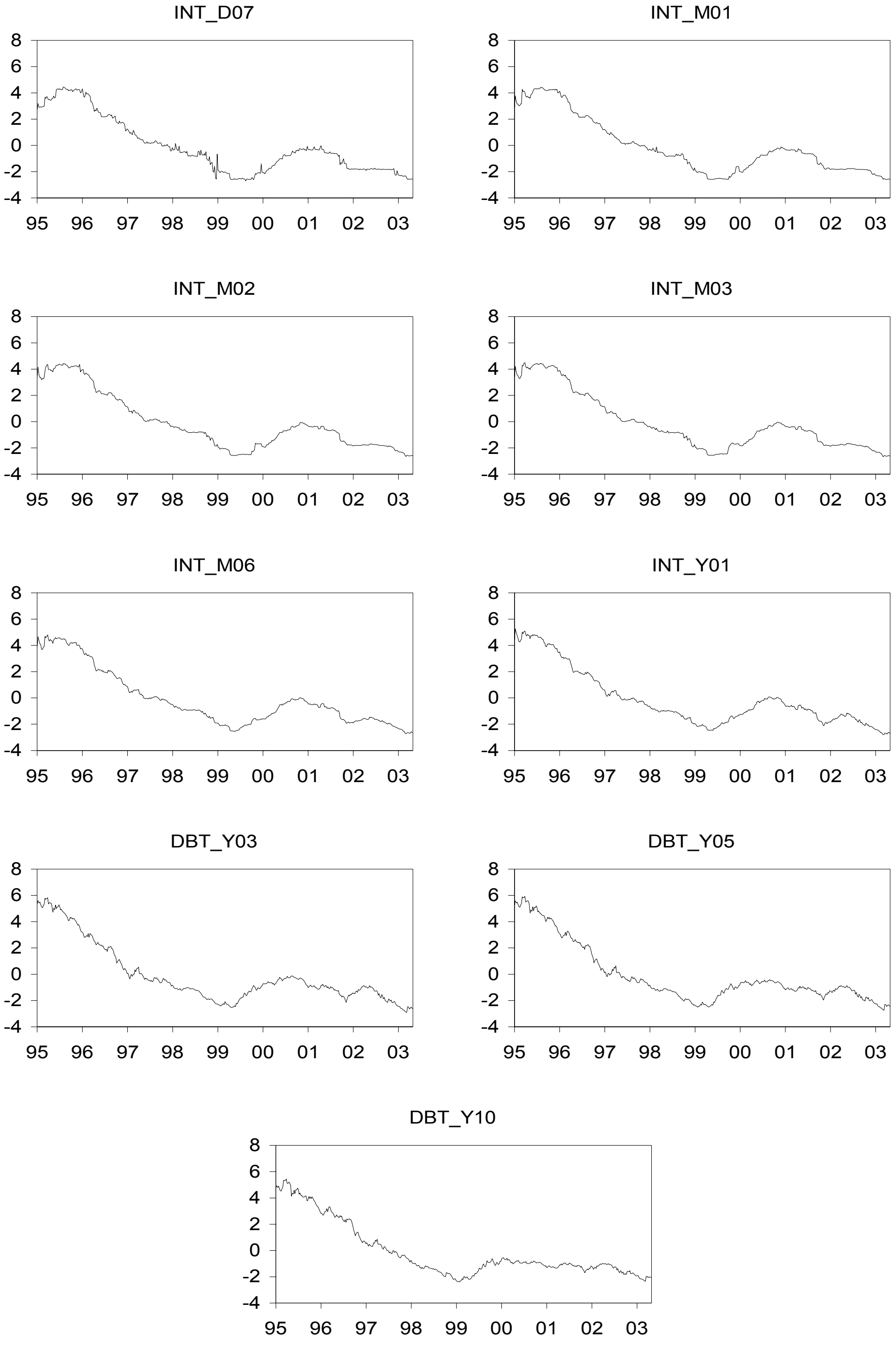

FIGURE 2
Interest rates in first differences.

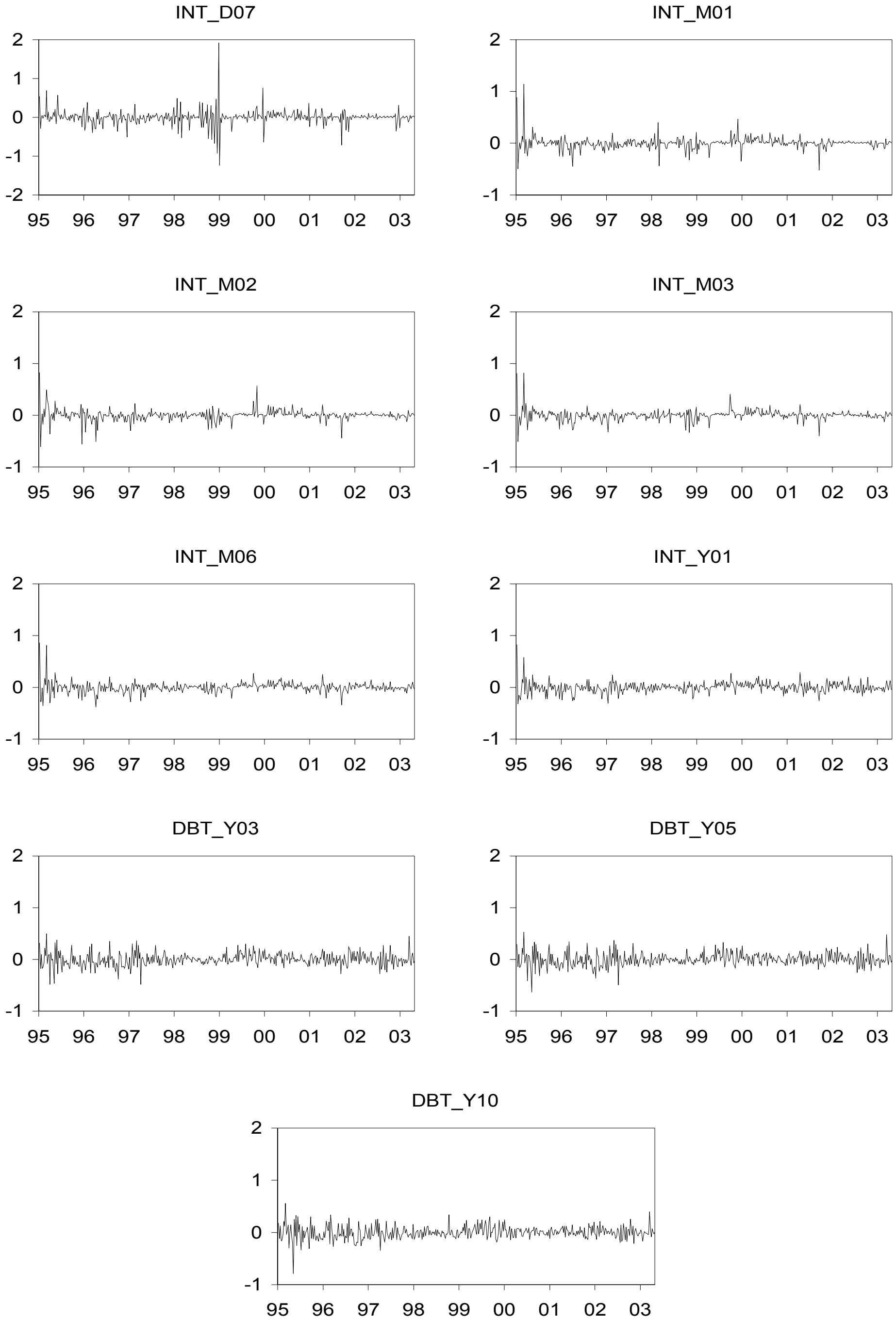

FIGURE 3
Interest rate spreads