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ARE THERE ASYMMETRIES IN THE RESPONSE OF BANK INTEREST RATES TO MONETARY SHOCKS?

by Leonardo Gambacorta[*] and Simonetta Iannotti[**]

Abstract

This paper examines the velocity and asymmetry of the response of bank interest rates to monetary policy shocks. Using an Asymmetric Vector Error Correction Model (AVECM), it analyses the pass-through of changes in money market rates to retail bank interest rates in Italy in the period 1985-2002. The main results of the paper are: 1) the speed of adjustment of bank interest rates to monetary policy changes increased significantly after the introduction of the 1993 Consolidated Law on Banking; 2) interest rate adjustment in response to positive and negative shocks is asymmetric in the short run, but not in the long run; 3) banks adjust their loan (deposit) rate faster during periods of monetary tightening (easing); 4) this asymmetry almost vanished since the 1990s.

JEL classification: E43, E44, E52.

Keywords: monetary policy transmission, interest rates, asymmetries, liberalization.

Contents

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1. Introduction[1]

This paper examines the velocity and asymmetry of the response of bank interest rates to monetary policy shock. These aspects are very important for understanding monetary transmission mechanisms: a change in the monetary stance is effective only if monetary impulses are transmitted quickly to other rates and if the new structure of interest rates affects real expenditure. Asymmetric behaviour of bank interest rates in the case of a monetary tightening or easing could have different effects on output and prices, and therefore knowing how much, how quickly and how symmetrically a change in the monetary interest rate is transmitted to bank rates is extremely important for the conduct of monetary policy. Moreover, an asymmetric response of banking rates also has major consequences for profit margins, interest rate risks and the overall performance of the banking industry.

The empirical literature so far has documented that lending and deposit rates respond sluggishly to money market rate changes.[2] The studies for Italy refer to the 1980s and early 1990s, before the enactment of the 1993 Consolidated Law on Banking which has fostered competition in the banking sector. One of the aims of this paper is to examine whether the increased competition has had any effect on interest rate setting: the financial liberalization process of Italy's banking industry in the 1990s should have led to a faster adjustment of bank interest rates to monetary policy changes compared with the 1980s, when a certain degree of stickiness in bank interest rates could be observed.

We analyze the simultaneous interactions between three bank rates (on current accounts, on short-term lending and on the interbank market) and the monetary policy indicator (the rate on repurchase agreements) in two separate periods. The first (1985:01-1993:08) coincides with the partial liberalization of the banking system, while in the second

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[2] Among cross-country studies, see Cottarelli and Kourelis (1994), Borio and Fritz (1995) and de Bondt et al. (2003). Among national studies see Cottarelli et al. (1995) and Angeloni et al. (1995) for Italy, Weth (2002) for Germany, and Berlin and Mester (1999) for the US.

period (1993:09-2002:12) the Banking Law was already in force. The paper tests for differences in the velocity of adjustment of banking rates to the monetary policy indicator and for the presence of asymmetric adjustments in the event of opposite monetary policy impulses (tightening or easing). The econometric framework used is an Asymmetric Vector Error Correction Model (AVECM) as in Lim (2001). The AVECM is based on a reformulation of the multivariate error correction model proposed by Johansen (1988; 1995), which allows for asymmetric behaviour both in the long and the short run. In particular, the model captures the interplay of long-run optimizing behaviour on the part of banks, embedded in the cointegration relationship, with their short-run adjustments, captured by the part in first difference.

The paper is organized as follows. Section 2 analyzes some institutional characteristics of the Italian banking sector. Section 3 gives a descriptive analysis of the data and identifies possible breaks in the estimation period. After an analysis of the characteristics of the VAR model in Section 4, Section 5 discusses the long-run relationship between the interest rates using Johansen's methodology. Section 6 presents the Asymmetric Vector Error Correction Model used to test for the presence of asymmetric behaviour depending on whether policy rates are increasing or decreasing. Model specification tests are reported in Section 7, while Section 8 contains the result of a simulation using the estimated AVECM for Italy. Robustness checks are presented in Section 9. The last section summarizes the main conclusions.

2. Some institutional characteristics of the Italian banking sector

Before discussing the econometric analysis of banks' interest rate setting, we briefly highlight the important measures to liberalize the markets and deregulate the intermediaries implemented over the last two decades (Ciocca, 2000). This institutional analysis will help us to identify the estimation periods with respect to different degrees of financial liberalization.

At the beginning of the 1980s the Italian banking system was quite tightly regulated: 1) foreign exchange controls were in place; 2) the establishment of new banks and the opening of new bank branches were subject to authorization;[3] 3) competition was curbed by mandatory maturity specialization, with special credit institutions operating at medium-long term maturities and commercial banks at short-term; 4) the quantity of bank lending was subject to a ceiling.

All these restrictions were gradually removed between the mid-1980s and the early 1990s (Cottarelli et al., 1995; Passacantando, 1996; Angelini and Cetorelli, 2002): 1) the ceiling on lending was abolished de facto in 1985; 2) foreign exchange controls were lifted between 1987 and 1990; 3) branching was liberalized in 1990; 4) the 1993 Consolidated Law on Banking allowed banks and special credit institutions to perform all banking activities.[4]

On the basis of these institutional characteristics of the Italian banking system, we divide the estimation period into two parts. The first sub-sample (1985:01-1993:08) refers to the period of partial liberalization. Previous periods are excluded because the presence of ceilings on lending could influence the results. The second sample (1993:09-2002:12) starts with the introduction of the Consolidated Law on Banking and refers to a period in which all restrictions were largely removed.

3. Data

The price setting behaviour of Italian banks is analyzed using four interest rates: the average rate on short-term lending (i_{L}), the average rate on current accounts (i_D), the three-month interbank rate (i_{B}) and the rate on repurchase agreements between the central bank and credit institutions (i_{M}).[5] A graphical analysis of the series is reported in Figure 1. It shows a high correlation between the series, suggesting the possibility that they are cointegrated.

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[3] Before 1987 the Bank of Italy authorized the opening of new branches on the basis of a 4-year plan reflecting estimated local needs for banking services.

[4] The 1993 Consolidated Law on Banking, introduced in September, completed the enactment of the institutional, operational and maturity despecialization of the Italian banking system and ensured the consistency of supervisory controls and intermediaries' range of operations with the single market framework. The business restrictions imposed by the 1936 Banking Law, which distinguished between banks that could raise short-term funds ("aziende di credito") and those that could not ("Istituti di credito speciale"), was eliminated. For more details see the Annual Report of the Bank of Italy for 1993.

[5] Data are available on the Internet site of the Bank of Italy (www.bancaditalia.it).

The choice of the rate on domestic short-term lending has two main advantages. First, it excludes credit directly channeled through legal requirements (i.e. lending to housing and rural sectors) and foreign exchange operations. Second, short-term loans are typically not collateralized and this allows the effects of the "balance sheet" channel to be isolated (Mishkin, 1995; Oliner and Rodebusch, 1996; Kashyap and Stein, 1997). Broadly speaking, the pass-through from market interest rates to the interest rate on loans does not depend on market price variations that influence the value of collateral. Nearly half of banks' business is done at this rate.

The deposit rate is the weighted average rate paid on current accounts, which are highly homogenous deposit products. Current accounts are the most common type of deposit (at the end of 2002 they represented around 70 per cent of total bank deposits and passive repos). Current accounts allow unlimited checking for depositors, who can close the account without notice. The bank, in turn, can change the interest paid on the account at any time.

Both bank rates are posted rates that are changed at discrete intervals (often less than weekly, see Green, 1998). In our case, the monthly frequency of the data is sufficient to capture all relevant changes due to monetary policy shocks. Both rates are before tax.

The interbank rate is included in the model because, especially in the first period of partial liberalization, the transmission of monetary policy impulses to the interbank rate could take more than a month (see, among others, Amisano et al., 1997).

The interest rate taken as monetary policy indicator is that on repurchase agreements between the Bank of Italy and credit institutions in the period 1985:01-1998:12, and the interest rates on main refinancing operations of the ECB in the period 1999:01-2002:12. As pointed out by Amisano et al. (1997) and Buttiglione and Ferri (1994), in the period under investigation the repo rate mostly affected the short-term end of the yield curve and it represented the value to which market rates and bank rates eventually tended to converge. It is worth noting that the interest rate on main refinancing operations of the ECB does not present any particular break with the repo rate at the beginning of stage three of EMU. The Augmented Dickey Fuller (*ADF*) tests provided in Table 1 clearly show that all the series are I(1) without drift.

The behaviour of bank interest rates in Italy reveals some stylized facts (see Figure 1). First, especially in the 1980s, the interest rate on current accounts was quite sticky to monetary policy changes;[6] this rigidity diminished after the introduction of the 1993 Banking Law. Second, there has been a considerable fall in average rates since the end of 1992.

The main reason for the fall in bank rates is probably the successful monetary policy adopted to reduce inflation in Italy in order to meet the Maastricht criteria and the third stage of EMU. As a result, the interbank rate decreased by more than 10 percentage points in the period 1993-1999. Excluding the 1995 episode of turbulence in the foreign exchange markets, it was only in the third quarter of 1999 that it started to climb, edging upwards until the end of 2000 and then declining thereafter. From a statistical point of view, this behaviour calls for the investigation of a possible structural break. Figure 2 shows sequential tests for changes in the mean of each bank interest rate. The hypothesis of this procedure, which was developed by Banerjee et al. (1992), is that if there is a break, its date is not known a priori but rather is gleaned from the data. The results clearly show that unit-root/no-break null can be rejected at the 2.5 per cent critical value level against the stationarity/mean-shift alternative for the period 1995:03-1998:09. In this period, which coincides with the convergence process towards stage three of EMU, it is necessary to investigate the existence of a shift in the mean in the long-run relationship between interest rates (see Section 6).

[6] Deposit interest rate rigidity in the 1980s has been extensively analyzed for the US as well. Among the market factors that have been found to affect the responsiveness of bank deposit rates are the direction of the change in market rates (Ausubel, 1992; Hannan and Berger, 1991), if the bank interest rate is above or below a target rate (Hutchison, 1995; Moore, Porter and Small, 1990; Neumark and Sharpe, 1992) and market concentration in the bank deposit market (Hannan and Berger, 1991). Rosen (2001) develops a model of price settings in the presence of heterogeneous customers explaining why bank deposit interest rates respond sluggishly to some extended movements in money market rates but not to others. Hutchison (1995) presents a model of bank deposit rates that includes a demand function for customers and predicts a linear (but less than one-to-one) relationship between market interest rate changes and bank interest rate changes. Green (1998) claims that the rigidity is due to the fact that bank interest rate management is based on a two-tier pricing system; banks offer accounts at market-related interest rates and at posted rates that are changed at discrete intervals.

4. The VAR model

The monetary transmission mechanism is explained using a four-variable VAR system: bank interest rates i_L, i_D and i_B are the endogenous variables that react to exogenous changes in the monetary policy indicator i_M in the two sub-periods.

The starting point of the multivariate analysis is the following reduced-form VAR:

$$\begin{aligned} y_t &= \mu + \sum_{k=1}^{p} \Phi_t y_{t-k} + \sum_{k=0}^{p} \Theta_t i_{M\,t-k} + \varepsilon_t \qquad t = 1,\dots,T \\ \varepsilon_t &\sim \mathrm{VWN}(0,\Sigma) \end{aligned} \tag{1}$$

where y_t=[i_L, i_D, i_B] and ε_t is a vector of white noise residuals. The deterministic part of the model includes a constant, while a trend is excluded a priori because there is nothing in economic theory to suggest that nominal interest rates should exhibit a deterministic time trend (Hamilton, 1994).[7]

In choosing the lag length of the *VAR* analysis *p*, several different criteria are used. The classical *LR* tests (with a small sample correction suggested by Sims, 1980) and the information criteria (Akaike and Schwarz) give evidence in favour of a model with 2 lags in the first sub-sample and 4 lags in the second sub-sample (see Table 2).

The analysis of the system shows serially uncorrelated residuals in both models. However, normality of the VAR is not achieved. The residual plot indicates that the non-normality could be attributable to few detected outliers.

A significant improvement in the stochastic properties of the VAR model for the first period is obtained by adding two dummies to capture the effects of monetary policy impulses in 1990 and 1992.[8] These dummies are in correspondence of specific monetary policy

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[7] The monetary policy interest rate has been considered an exogenous variable. This hypothesis has been tested in a VAR model where all interest rates are treated as endogenous variables. The null hypothesis of weak exogeneity of the monetary policy indicator has been accepted with a p-value of 20.5 per cent. Following Harris (1995), we have therefore removed the equation for the monetary policy indicator from the system.

[8] The first one *du90*, reflects Bank of Italy interventions soon after capital movement liberalization (May 1990). "In June, to prevent liquidity conditions from becoming excessively tight, the Bank of Italy made gross temporary purchases of securities in the secondary market totaling 21 trillion lire". In September, the market was not attracted by medium-term securities. "With the aim of redirecting demand towards the longer end of the market, the Bank of Italy supplied only a very small quantity of these instruments for a short period lasting

interventions. As regards the second sub-period, one point dummy in 1995:03 is necessary to take account of the spikes in interest rates due to turbulence in the foreign exchange markets.

Even if the stochastic properties of the model improve significantly, reaching a normal residual distribution (see Table 3), the inclusion of these dummies affects the underlying distribution on which cointegration tests depend, so that the critical values reported thereafter are only indicative.

5. Cointegration properties of interest rates

Cointegration can be analyzed by re-expressing equation (1) as a reduced-form error correction model:

$$\begin{aligned} &\Delta y_t = \Pi(\mu, y_{t-1}, i_{M\,t-1}) + \sum_{k=1}^{p-1} \Gamma_t \Delta y_{t-k} + \sum_{k=1}^{p-1} \Psi_t \Delta i_{M\,t-k} + \eta\Phi_t + \varepsilon_t \qquad t = 1,\dots,T \\ &\Pi = \alpha\beta' \end{aligned} \tag{2}$$

where Φ represents a vector of dummies. The constant is included in the cointegration space; in fact, theory suggests that the constant captures the possible existence of mark-up or mark-down in the long-run relationship between interest rates. In the second sub-period, given the structural break in the mean, the convergence dummy[9] as well as the constant are allowed to lie in the cointegration space.

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until the middle of the month. This caused liquidity to become abundant and banks' excess reserves averaged around 8 trillion lire in the first two ten-day periods of September. The *REPO* rate fell to 6.7 per cent". "In October there was a net foreign exchange outflow of 2.3 trillion lire despite the placement of a 1 billion ecu bond issue abroad. The central bank counteracted a substantial creation of liquidity through the Treasury current account by making temporary security sales of 13.1 trillion lire at rates of around 11per cent which were appreciably higher than the rates prevailing in September" (see Bank of Italy, "Economic Bullettin", October 1986, pp. 37-39). *Du90* was set to +1 in 90:6 and 90:9 and to -1 in 90:10. The second dummy, *du92*, reflects central bank operations during the 1992 currency crisis. After two increases in the official discount rate (from 12 to 13.75 per cent in July and to 15 per cent in September) in order to maintain the ERM parity monetary conditions were relaxed in November (from 15 to 13 per cent) after Italy left the ERM. In order to capture monetary policy behaviour, dum92 has a -1 on 92:7 and 92:9 a +1 on 92:11. It is worth noting that *du90* and *du92* gave a better result than using five point dummies (one for each date discussed above) coupled with a considerable gain in efficiency.

[9] The convergence dummy (dum) takes the value 1 between 1995:03 and 1998:09 and zero elsewhere. It represents a monetary policy stance geared to achieve the convergence of Italian interest rates towards those prevailing in the euro area.

This framework can be used to apply Johansen's trace test to verify the order of integration of the matrix Π. In fact, the rank of Π determines the number of cointegrating vectors (r) such that α is a $n \times r$ matrix of loading coefficients and β is a $n \times r$ matrix of cointegrating vectors.

The results are reported in Table 4. Johansen's cointegration rank statistics show the presence of 3 cointegrating relationships in the model. The hypothesis of the existence of three cointegrating vectors is consistent with a strong a-priori economic view because if we consider a set of nominal interest rates, the non–stationary driving force for all of them is likely to be the inflationary process. Nevertheless, as discussed in the previous section, the presence of dummy variables in the model affects the underlying distribution on which Johansen's cointegration test depends, so that the critical values reported in the first part of Table 4 are only indicative. Therefore, in order to provide the robustness of the rank result, a Hansen and Johansen (1993) iterative procedure is investigated. The outcome, presented in Figure 3, suggests that the evidence of rank 3 is strongly consistent.

As for the economic interpretation of the cointegrating relationship, we suppose that the interbank rate is equal to the exogenous monetary policy rate plus a mark-up, μ_B. The latter is equal to zero if interbank lending is considered a risk-free activity.

$$i_B = i_M + \mu_B \tag{3}$$

Economic theory on oligopolistic (and perfect) competition suggests that, in the long run, both bank rates (on lending and deposits) should be related to the interbank rate that represents the cost of banks' refinancing. For example, Freixas and Rochet (1997) show that in a model of imperfect competition among N banks, the relationships between the three interest rates become:

$$i_L = i_B + \mu_L \tag{4}$$

$$i_D = (1 - \xi) i_B + \mu_D \tag{5}$$

where $\mu_L = \gamma_L + i_L'(L^*)\frac{L^*}{N}$ and $\mu_D = \gamma_D + i_D'(D^*)\frac{D^*}{N}$ are constants. In the unique Cournot equilibrium each bank *x* sets the same quantity of loans ($L_i = \frac{L^*}{N}$) and deposits ($D_i = \frac{D^*}{N}$). A part of deposits (ξ) is invested in compulsory (or free) reserves. The mark-up μ_L (mark-down μ_D) is influenced by the constant marginal cost of intermediation on lending γ_L (deposits γ_D) and by the elasticity of the loan (deposit) demand function evaluated at the optimum. It is worth noting that in the case of perfect competition ($N \to \infty$), the last part in μ_L and μ_D goes to zero and long-run relationships are independent of loan and deposit demand functions.

The normalized cointegrating relationships are presented in the second part of Table 4, with the associated standard errors. No identification restrictions are imposed on the cointegrating space. The results give us the following insights. First, in the period of partial liberalization both the long-run elasticities of bank rates with respect to the monetary policy indicator and the loading coefficients have lower absolute values. This result leads to an increase in competition in the 1990s. In particular, the lower values of the loadings in the first part of the sample period indicate a more sluggish reversion to the long-run equilibrium in the case of an exogenous shock.

Second, given that some of the loading coefficients are not statistically different from zero in both periods, the model can be simplified. In particular, the third cointegrating relationship does not enter the first two equations for i_L and i_D (α_{D3} and α_{L3} are statistically not different from zero), while the first two cointegrating relationships do not enter the equation for i_B (α_{B1} and α_{B2} are zero). This means that the interbank rate is weakly exogenous with respect to the two banking rates and that it responds directly to the monetary policy indicator. Indeed, exogenous shocks in the long-run relationships that drive both bank rates do not influence the interbank rate. This result is consistent with a causal chain between interest rates of the type: $i_M \to i_B \to (i_D, i_L)$. The null hypothesis $\alpha_{D3}=\alpha_{L3}=\alpha_{B1}=\alpha_{B2}=0$ is accepted for both sub-samples with p-values of, respectively, 0.24 and 0.10 per cent. As we will see in the next section this result will be used to reduce the number of parameters in the asymmetric model.

6. The Asymmetric Vector Error Correction Model (AVECM)

The model analyzed so far is symmetric. However, interest rate adjustments may be asymmetric in size and speed. For example, in the case of a monetary tightening, if banks had some market power, they could increase their loan rate by more and faster than their deposit rate, and vice versa, in the case of an easy monetary policy. This behaviour implies asymmetric adjustment of bank rates both in magnitude and speed, and therefore the multivariate framework described by (2) should be extended to allow for asymmetric behaviour in the long-run cointegrating relationship (β), the loading coefficients (α) and lagged responses of variables in delta (Γ).

Following Lim (2001), the VECM system may be expanded to allow for asymmetric adjustments in both long-run and short-run behaviour. Preliminarily we test whether it is worthwhile to pass from the symmetric model to the more general asymmetric model (Teräsvirta, Tjøstheim and Granger, 1994). This test is particularly useful because if no significant gain is detected using the more general model, it is possible to stop further investigation. Tests for the two periods give as results: $\chi^2(56.9, 30)=0.00$ and $\chi^2(101.9, 54)=0.00$, which confirm the need for an asymmetric approach to the problem.

In order to reduce the number of parameters to be estimated, in the asymmetric model we then use the result $\alpha_{D3}=\alpha_{L3}=\alpha_{B1}=\alpha_{B2}=0$ which is valid in both sub-periods. This helps us to increase the number of degrees of freedom, especially in the second period.

The VECM system (2) with three cointegrating vectors can be reformulated as:

$$\begin{aligned}
\Delta i_{D,t} &= (\alpha_{D1} + \alpha^*_{D1} d_t)[i_{D,t-1} - (\mu_D + \mu^*_D d_t) - (\beta_D + \beta^*_D d_t)\, i_{B,t-1}] + \\
&\quad + (\alpha_{D2} + \alpha^*_{D2} d_t)[i_{L,t-1} - (\mu_L + \mu^*_L d_t) - (\beta_L + \beta^*_L d_t)\, i_{B,t-1}] + \\
&\quad + \sum_{k=1}^{p-1} (\delta_{Dk} + \delta^*_{Dk} d_t)\Delta i_{D,t-k} + \sum_{i=1}^{p-1} (\varphi_{Dk} + \varphi^*_{Dk} d_t)\Delta i_{L,t-k} + \\
&\quad + \sum_{k=1}^{p-1} (\psi_{Dk} + \psi^*_{Dk} d_t)\Delta i_{B,t-k} + \sum_{k=0}^{p-1} (\phi_{Dk} + \phi^*_{Dk} d_t)\Delta i_{M,t-k} + \Gamma_D \Phi_t + \varepsilon^L_t
\end{aligned} \tag{6}$$

$$\begin{aligned}\Delta i_{L,t} &= (\alpha_{L1} + \alpha^*_{L1} d_t)[i_{D,t-1} - (\mu_D + \mu^*_D d_t) - (\beta_D + \beta^*_D d_t)\, i_{B,t-1}] + \\ &+ (\alpha_{L2} + \alpha^*_{L2} d_t)[i_{L,t-1} - (\mu_L + \mu^*_L d_t) - (\beta_L + \beta^*_L d_t)\, i_{B,t-1}] + \\ &+ \sum_{k=1}^{p-1} (\delta_{Lk} + \delta^*_{Lk} d_t)\Delta i_{D,t-k} + \sum_{k=1}^{p-1} (\varphi_{Lk} + \varphi^*_{Lk} d_t)\Delta i_{L,t-k} + \\ &+ \sum_{k=1}^{p-1} (\psi_{Lk} + \psi^*_{Lk} d_t)\Delta i_{B,t-k} + \sum_{k=0}^{p-1} (\phi_{Lk} + \phi^*_{Lk} d_t)\Delta i_{M,t-k} + \Gamma_L \Phi_t + \varepsilon^L_t \end{aligned} \tag{7}$$

$$\begin{aligned}\Delta i_{B,t} &= (\alpha_{B3} + \alpha^*_{B3} d_t)[i_{B,t-1} - (\mu_B + \mu^*_B d_t) - (\beta_B + \beta^*_B d_t)\, i_{M,t-1}] + \\ &+ \sum_{k=1}^{p-1} (\delta_{Bk} + \delta^*_{Bk} d_t)\Delta i_{D,t-k} + \sum_{k=1}^{p-1} (\varphi_{Bk} + \varphi^*_{Bk} d_t)\Delta i_{L,t-k} + \\ &+ \sum_{k=1}^{p-1} (\psi_{Bk} + \psi^*_{Bk} d_t)\Delta i_{B,t-k} + \sum_{k=0}^{p-1} (\phi_{Bk} + \phi^*_{Bk} d_t)\Delta i_{M,t-k} + \Gamma_B \Phi_t + \varepsilon^B_t \end{aligned} \tag{8}$$

In this AVECM, the three cointegrating vectors are normalized on rates i_D, i_L and i_B. The constant terms μ_D and μ_L are the intermediation margins; μ_B is the mark-up between i_B and i_M; β_D and β_L represent the long-run elasticities of i_D and i_L with respect to the interbank rate; β_B is the elasticity between the interbank rate and the monetary policy indicator. The loading coefficients are represented by α_{kr}, with k and r indicating, respectively, the equation (k=D,L,B) and the number of the cointegrating relationship (r=1,2,3). The parameters δ, φ, ψ, and ϕ specify the lagged coefficients.

Parameters that refer to asymmetric behaviour are those with the superscript "*". These are interacted with the dummy variable *d,* which captures the differential effects of increases and decreases in the monetary policy indicator. There are two possible stances of monetary policy: monetary loosening (a negative change in the repo rate) and monetary tightening (a positive change in the repo rate). Therefore *d* is defined according to the following scheme:

$$d = \begin{cases} 1 & \text{if } \Delta i_{\mathrm{M}} < 0 \\ 0 & \text{if } \Delta i_{\mathrm{M}} > 0 \end{cases}$$

In a few cases no monthly changes are detected in the monetary indicator (Δi_M=0). In these months, a monetary easing (tightening) is considered, d=1(d=0), if the interbank interest rate shows a reduction (increase), leading to easier (more difficult) access to interbank liquidity. Figure 4 shows the changes in the monetary policy indicator in the two periods.

7. Testing asymmetry and the reduced-form model

Starting from the model described in equations (6)-(8), we follow a general to specific strategy to test for asymmetry. Nevertheless, this approach is not interpreted as a mechanical reduction process that implies dropping all insignificant parameters (Pagan, 1990). The removal of every insignificant parameter is done to control for the multivariate significance level of the model. All tests for asymmetry are reported in Table 5.

Asymmetry is tested considering the null hypothesis of zero restrictions on the dummy variables. The test for asymmetry in the loading coefficients (see part A of Table 5) supports the hypothesis of a different adjustment to disequilibrium gaps. On closer analysis, the asymmetry in the first period is contained in the loading coefficients α_{D1}^* and α_{L2}^*; in this case the model can be further simplified because α_{D2} and α_{L1} are statistically not different from zero.[10] In the second period a single asymmetry is detected for α_{L1}^*.

From an economic perspective this means that in the first period there is a greater difference in the velocity of adjustment of both bank rates towards the long-run equilibrium. In particular, the signs of the coefficients show that the interest rate on deposits responds faster when the deviation from the long-run equilibrium is caused by an easy monetary policy. On the contrary, the adjustment towards the long-run equilibrium is faster for lending rates in the case of monetary tightening. In the second period, the only asymmetry is detected in the response of the lending rate to a deviation in the long run relationship between the interest rate on deposits and the interbank rate. This result suggests that after the introduction of the Banking Law there was a greater interplay between lending and deposits price strategy.

By contrast, the test for asymmetry in the intercept and elasticities of the long-run relationships (parts B and C in Table 5) always fails to support the hypothesis of a different

[10] The likelihood ratio for $\alpha_{D2}=\alpha_{D2}^*=\alpha_{L1}=\alpha_{L1}^*=0$ is given by χ^2 (4)=5.57 with a p-value of 0.23 per cent.

equilibrium due to the characteristics of the monetary policy impulse. This means that in the long run the pass-through from money market rates to bank rates has the same size independent of the sign of the shock. This is also consistent with the idea that in the long run the equilibrium between interest rates is unique.

The significance of asymmetries in the lagged terms (see part D of table 5) gives information about the dynamic path of adjustment in the short run. The results show that in both sub-samples bank interest rates react asymmetrically to short-term changes in the monetary policy indicator. However, the asymmetric effect on the interbank rate is statistically significant only in the first period, and vanishes in the 1990s. There is also an asymmetry in the autoregressive part of the equation for loan interest rates for the second period; however, since the sign of the asymmetric coefficient is positive it tends to counterbalance the asymmetric effects detected in the case of a change in the monetary policy indicator. In other words, when a monetary easing occurs, the reduction of the short-term interest rate in the first months is counterbalanced by a positive autoregressive coefficient, which reflects an increase in the velocity of adjustment towards the new equilibrium. Table 6 presents the results for the reduced trivariate system, including the significant asymmetric short-term effects.

8. A simulation: adjustment to positive and negative shocks

In order to evaluate the effect of an exogenous monetary policy shock a simulation exercise is performed to generate time paths for the three bank rates. Figure 5 presents the adjustment paths of the three bank interest rates to positive and negative changes in i_M. In particular, the policy experiment consists in increasing (decreasing) the repo rate by one per cent, starting from a base when the trivariate system is in equilibrium. To make simulations more graphically comparable, the effects for easy monetary policy have been multiplied by -1. The main results are the following.

In both regimes (partial liberalization and complete deregulation), the cumulative long-run changes in i_L, i_D and i_B to a monetary shock are symmetric. Consistently with economic theory, the hypothesis of a long-run unitary elasticity between both the short-term lending rate and the interbank rate and the monetary policy indicator is largely accepted (β_L=β_B=1).[11] On the contrary, given the presence of the reserve coefficient, the elasticity between i_D and i_M, β_D, is around 0.7 in both sub-periods. This is consistent with the work of Cottarelli et al. (1995) for the first period and Gambacorta (2005) for the second period.

In the short run, lending rates adjust faster with rising interest rates and less markedly when interest rates are falling. On the contrary, interest rates on deposits tend to converge more rapidly to the long-run equilibrium in the case of a monetary easing and more sluggishly in the case of a tightening. However, these differences in short-run adjustments diminish over time. Indeed, after a year the gap between the response of the lending rate to a tight and an easy monetary policy is 14 basis points in the first period and 3 basis points in the second. In the case of deposits the difference is -12 basis points in the period of partial liberalization and vanishes in the deregulation period.

Asymmetries in the adjustment of the interbank rate to a monetary shock are detected only in the first period. This reflects the existence, from the beginning of the 1990s, of an efficient screen-based market for interbank deposits (Mercato Interbancario dei Depositi, MID) in Italy, which led to a reduction in the number of bilateral current accounts between banks and a strong increase in competitiveness.

9. Robustness checks

The robustness of the results is checked in several ways. The first test is to estimate a simple model for the second sub-sample (1993:09-2002:12) which excludes the interbank rate. This specification helps us to check if the results on the asymmetric effects in Section 7 are robust in terms of a greater number of degrees of freedom. The exclusion of the interbank rate from the model can be adopted because of the high velocity of adjustment of the interbank rate to the policy rate in the 1990s. However, a formal test of exclusion of this variable from the model can be accepted only marginally (the p-value is 0.011). The simplified version of the model turns out to have the same characteristics as the more general

[11] As for the test β_L=β_B=1, the likelihood ratio test statistic is 0.981 with a p-value of 0.612 in the first period and 0.402 with a p-value of 0.818 in the second period.

one (the same lags and cointegration properties). As for asymmetries, the only difference is the lack of significance of α^*_{L1} (the null hypothesis can be accepted with a p-value of 0.125). The simulation exercises confirm the existence of a difference in adjustment only in the short-run (the gaps between the response of bank rates to a tight and an easy monetary policy vanish after eight months). No asymmetry is detected in the long run.

The second test considers whether different fiscal treatments over the sample period could have changed deposit demand (from June 1996 the interest rate on deposits is subject to 27 per cent tax, deducted at source; 12.5 per cent before). However, when the net interest rate on current accounts is used in place of the gross rate, nothing changes.

The third robustness check analyzes the cointegration properties in a model where all variables are endogenous. Even in this case the λ-trace test shows the existence of three cointegrating relationships. Loading coefficients and long-run elasticities remain the same.

10. Conclusions

In this paper we investigate velocity and asymmetry in the response of bank interest rates to monetary policy shocks in Italy in the period 1985-2002. Understanding asymmetric responses of bank rates is important in two respects: first, because of their potential impact on output and prices, and second, because it gives insights into banks' behaviour in their relationship with customers and more generally regarding the evolution of competitive conditions in the credit and deposit markets. Using an Asymmetric Vector Correction Model (AVECM) that allows for different behaviour in both the long-run and the short-run, we obtain the following results: 1) the speed of adjustment of bank interest rates to monetary policy changes increased significantly after the introduction of the 1993 Banking Law; 2) interest rate adjustments, in response to positive and negative shocks, are asymmetric in the short run, but not in the long run, consistently with the idea that in the long run the equilibrium is unique; 3) banks adjust their loan (deposit) prices at a faster rate during periods of monetary tightening (easing); 4) this asymmetry has almost vanished since the 1990s.

Tables and figures

Table 1

AUGMENTED DICKEY-FULLER TESTS

Only two models are compared: Model B, which includes a constant and Model A, which has no deterministic components. Following Hamilton (1994), the presence of a linear trend in interest rates is not considered because it is theoretically inconsistent. Starting from the more general model the progressive tests establish what is the most appropriate deterministic component (Model B → Model A). The lagged differences in the models are included to obtain white noise residuals. The length of the lag (p*) used in the model for each series is chosen by comparing three different information criteria (Schwarz, Hannan-Quinn and Final Prediction Error), taking into consideration the necessary condition that the errors should be white noise. Further information on the test is provided by Said and Dickey (1984) and Dolado and Jenkinson (1987).

Series	p*	Model B $\Delta y_t = \mu + \rho\, y_{t-1} + \sum_{j=1}^{p^*-1} \rho_j \Delta y_{t-1} + \varepsilon_t$			Model A $\Delta y_t = \rho\, y_{t-1} + \sum_{j=1}^{p^*-1} \rho_j \Delta y_{t-1} + \varepsilon_t$
		ρ=0 τ_μ	μ=0 $\tau_{\alpha\mu}$	ρ=μ=0 Φ_1	ρ=0 τ
			I:1985:01-1993:08		
i_L	3	-2.66	2.45	3.96	-1.06
i_D	4	-2.81	2.49	4.53	-2.11
i_B	4	-2.73	2.63	4.10	-0.99
i_M	6	-2.74	2.67	3.89	-0.73
			II: 1993:09-2002:12		
i_L	3	-0.77	0.40	0.85	-1.19
i_D	4	-1.09	0.25	1.53	-1.20
i_B	4	-0.79	0.31	0.86	-1.32
i_M	2	-0.66	-0.01	1.40	-1.51
Critical values	5%	-2.89	2.54	4.71	-1.94
	1%	-3.51	3.22	6.70	-2.59

Table 2

LAG ORDER DETERMINATION

Information criteria: AK=Akaike and SC=Schwarz. The Likelihood Ratio (LR) is computed taking into account the small sample correction suggested by Sims (1980). GODF=Godfrey portmanteau test for autocorrelation of order 1.

Lag (h)	LR: L[h] vs L[h-1]	df	p-va	AK	SC	GODF p-va	LR: L[h] vs L[h-1]	df	p-va	AK	SC	GODF p-va
			I:1985:01-1993:08						II: 1993:09-2002:12			
1	n.a.	12	n.a.	-9.56	-8.29	0.000	n.a.	12	n.a.	-13.76	**-13.15**	0.239
2	72.323	12	**0.000**	-10.15	**-9.20**	**0.249**	30.987	12	0.002	-13.86	-12.96	0.506
3	41.333	12	0.051	**-10.61**	-9.14	0.122	29.007	12	0.004	-13.96	-12.75	0.000
4	25.982	12	0.101	-10.50	-8.92	0.071	39.957	12	**0.000**	**-14.20**	-12.68	**0.351**
5	9.658	12	0.646	-10.39	-8.48	0.001	9.462	12	0.663	-14.08	-12.28	0.100
6	19.349	12	0.080	-10.41	-8.20	0.003	10.590	12	0.564	-13.99	-11.88	0.309

Table 3

JARQUE-BERA NORMALITY TESTS

Normality is accepted when the p-value is larger than 5 per cent.

EQUATION	SKEWNESS	p-value	KURTOSIS	p-value	SKEW.&KURT.	p-value
			I: 1985:01-1993:08			
i_D	2.846	0.092	0.022	0.882	2.868	0.238
i_L	0.964	0.326	5.122	0.024	6.085	0.048
i_M	1.171	0.279	0.523	0.470	1.694	0.429
SYSTEM	4.137	0.247	3.567	0.312	7.704	0.261
			II: 1993:9-2002:12			
i_D	0.755	0.385	0.479	0.489	1.235	0.539
i_L	0.930	0.335	0.134	0.714	1.064	0.587
i_M	0.002	0.968	0.023	0.879	0.025	0.988
SYSTEM	1.635	0.652	1.707	0.635	3.342	0.765

Table 4

COINTEGRATION ANALYSIS

Test for the cointegration rank of the models in the two sub-samples. ** denotes rejection at the 1 per cent significance level; * denotes rejection at the 5 per cent significance level. The model for the second sub-sample also includes the convergence dummy in the cointegrating space. Johansen λ-trace tests take into account the adjustment for degrees of freedom proposed by Reimers (1993) for small samples. Asymptotic critical values are provided in Osterwald-Lenum (1992), although due to the presence of dummy variables they are only indicative.

	H_0: r=0	H_0: $r\leq 1$	H_0: $r\leq 2$
I: 1985:1-1993:08	57.35**	24.25*	11.11*
II: 1993:9-2002:12	68.61**	37.88**	11.57*

Cointegrating vectors

(standard errors in brackets)

I: 1985:1-1993:08

$$i_D = \underset{(0.136)}{0.568}\, i_B - \underset{(1.744)}{0.745}$$

$$i_L = \underset{(0.088)}{0.856}\, i_B + \underset{(1.126)}{3.634}$$

$$i_B = \underset{(0.096)}{0.892}\, i_M + \underset{(1.189)}{1.473}$$

II: 1993:9-2002:12

$$i_D = \underset{(0.018)}{0.691}\, i_B - \underset{(0.098)}{1.153} - \underset{(0.095)}{0.314}\, dumco$$

$$i_L = \underset{(0.027)}{1.039}\, i_B + \underset{(0.148)}{2.028} - \underset{(0.144)}{0.387}\, dumco$$

$$i_B = \underset{(0.073)}{1.010}\, i_M + \underset{(0.386)}{0.088} - \underset{(0.390)}{0.984}\, dumco$$

Loading coefficients

(standard errors in brackets)

I: 1985:1-1993:08

$$\alpha = \begin{pmatrix} \underset{(0.016)}{-0.047} & \underset{(0.030)}{0.000} & \underset{(0.016)}{-0.004} \\ \underset{(0.022)}{0.027} & \underset{(0.027)}{-0.096} & \underset{(0.023)}{-0.031} \\ \underset{(0.097)}{0.046} & \underset{(0.119)}{-0.008} & \underset{(0.063)}{-0.237} \end{pmatrix}$$

II: 1993:9-2002:12

$$\alpha = \begin{pmatrix} \underset{(0.086)}{-0.339} & \underset{(0.065)}{0.164} & \underset{(0.016)}{-0.016} \\ \underset{(0.066)}{0.168} & \underset{(0.050)}{-0.193} & \underset{(0.059)}{-0.003} \\ \underset{(0.133)}{0.034} & \underset{(0.788)}{0.146} & \underset{(0.012)}{-0.146} \end{pmatrix}$$

TESTS FOR ASYMMETRY

P-values. The symbols ***, ** and * represent significance at the 1, 5 and 10 per cent level. In the first sub-sample (1985:01-1993:08) the optimal p is equal to 2. In the second sub-sample (1993:09-2002:12) the optimal p is 4.

	(I) 1985:01-1993:08	(II) 1993:09-2002:12
A. Testing asymmetry in the loading coefficients		
$\alpha^*_{D1}=0$	0.041**	0.273
$\alpha^*_{D2}=0$	0.940	0.256
$\alpha^*_{L1}=0$	0.112	0.045**
$\alpha^*_{L2}=0$	0.021**	0.396
$\alpha^*_{B3}=0$	0.218	0.397
B. Testing asymmetry in the intercept of the long-run relationship		
$\mu^*_{D}=0$	0.528	0.903
$\mu^*_{L}=0$	0.181	0.933
$\mu^*_{B}=0$	0.517	0.621
C. Testing asymmetry in the elasticity of the long-run relationship		
$\beta^*_{D}=0$	0.576	0.584
$\beta^*_{L}=0$	0.219	0.252
$\beta^*_{B}=0$	0.555	0.858
D. Testing asymmetry in the short-term terms		
$\delta^*_{D1}=\ldots=\delta^*_{D\,p-1}=0$	0.760	0.254
$\varphi^*_{D1}=\ldots=\varphi^*_{D\,p-1}=0$	0.156	0.564
$\psi^*_{D1}=\ldots=\psi^*_{D\,p-1}=0$	0.896	0.041**
$\phi^*_{D1}=\ldots=\phi^*_{D\,p-1}=0$	0.076*	0.292
$\delta^*_{L1}=\ldots=\delta^*_{L\,p-1}=0$	0.703	0.402
$\varphi^*_{L1}=\ldots=\varphi^*_{L\,p-1}=0$	0.632	0.002***
$\psi^*_{L1}=\ldots=\psi^*_{L\,p-1}=0$	0.853	0.034**
$\phi^*_{L1}=\ldots=\phi^*_{L\,p-1}=0$	0.393	0.011**
$\delta^*_{B1}=\ldots=\delta^*_{B\,p-1}=0$	0.641	0.161
$\varphi^*_{B1}=\ldots=\varphi^*_{B\,p-1}=0$	0.232	0.515
$\psi^*_{B1}=\ldots=\psi^*_{B\,p-1}=0$	0.320	0.714
$\phi^*_{B1}=\ldots=\phi^*_{B\,p-1}=0$	0.043**	0.678

Table 6

COEFFICIENTS OF THE TRIVARIATE MODEL IN REDUCED FORM

Standard errors in brackets. The symbols ***, ** and * represent significance at the 1, 5 and 10 per cent level. Coefficients and standard errors for the dummies are not reported.

Equations for:	(I) 1985:01-1993:08			(II) 1993:09-2002:12		
	r_D	r_L	r_B	r_D	r_L	r_B
α_1	-0.044**			-0.253**	0.015	
	(0.018)			(0.093)	(0.085)	
α_2		-0.115***		0.037	-0.184***	
		(0.029)		(0.060)	(0.048)	
α_3			-0.187***			
			(0.035)			
α^*_1	-0.037**				0.143**	
	(0.018)				(0.034)	
α^*_2		0.047**				
		(0.022)				
μ_D	-1.470	-1.470	-1.470	-0.980***	-0.980***	-0.980***
	(1.850)	(1.850)	(1.850)	(0.101)	(0.101)	(0.101)
μ_L	1.838***	1.838***	1.838***	2.206***	2.206***	2.206***
	(0.145)	(0.145)	(0.145)	(0.060)	(0.060)	(0.060)
β_D	0.699***	0.699***	0.699***	0.700***	0.700***	0.700***
	(0.148)	(0.148)	(0.148)	(0.016)	(0.016)	(0.016)
β_L	1.000	1.000	1.000	1.000	1.000	1.000
	(-)	(-)	(-)	(-)	(-)	(-)
δ_1	0.174*			0.075	0.165**	
	(0.090)			(0.129)	(0.073)	
δ_2				-0.163**	-0.189***	
				(0.079)	(0.069)	
ϕ_0	0.020	0.063**	0.436***	0.037	0.214***	0.707***
	(0.015)	(0.018)	(0.055)	(0.034)	(0.063)	(0.709)
ϕ_1		0.040**	0.151***	-0.001		
		(0.017)	(0.045)	(0.042)		
ϕ_3					0.080	
					(0.059)	
ϕ^*_0				0.056*	-0.137**	
				(0.032)	(0.068)	
ϕ^*_1			-0.132**	0.066*		
			(0.055)	(0.037)		
ϕ^*_3					-0.124**	
					(0.062)	
φ_1	-0.029	0.429***		0.0354*	-0.159	
	(0.055)	(0.048)			(0.126)	
φ_2					0.011*	
					(0.006)	
φ_3				-0.032*		
				(0.019)		
φ^*_1					0.452***	
					(0.146)	
ψ_1	0.097***	0.096***			0.139***	
	(0.019)	(0.027)			(0.025)	
ψ_3					0.063*	
					(0.037)	
ψ^*_3					0.012**	
					(0.006)	

Figure 1

Interest rates

(monthly data, percentage points)

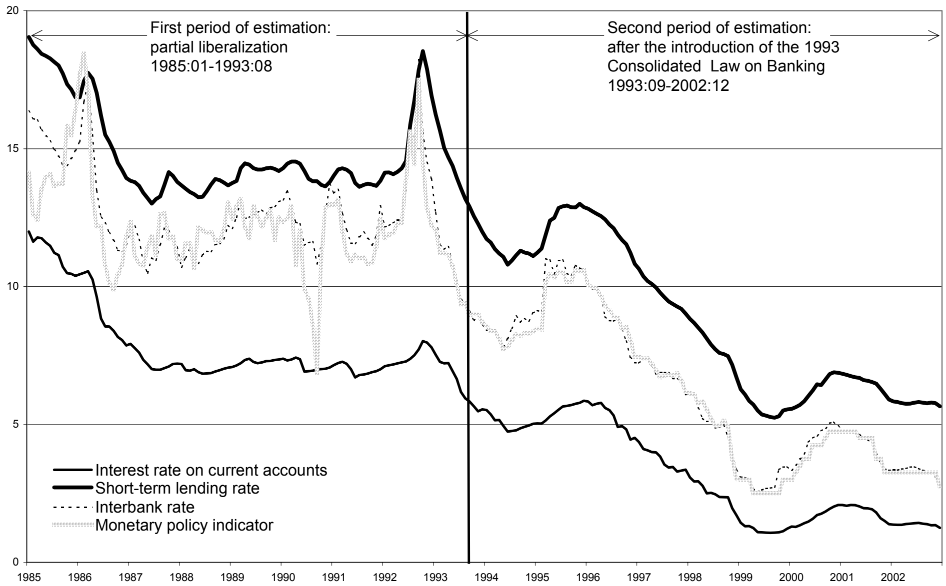
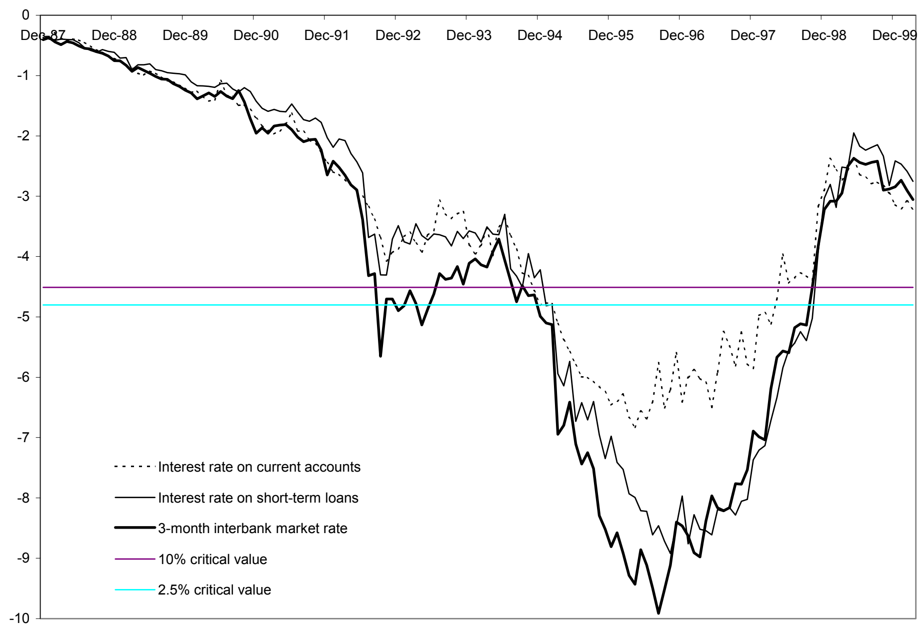


Figure 2

Search for mean shift breaks

(monthly data, sequential minimum unit root tests)

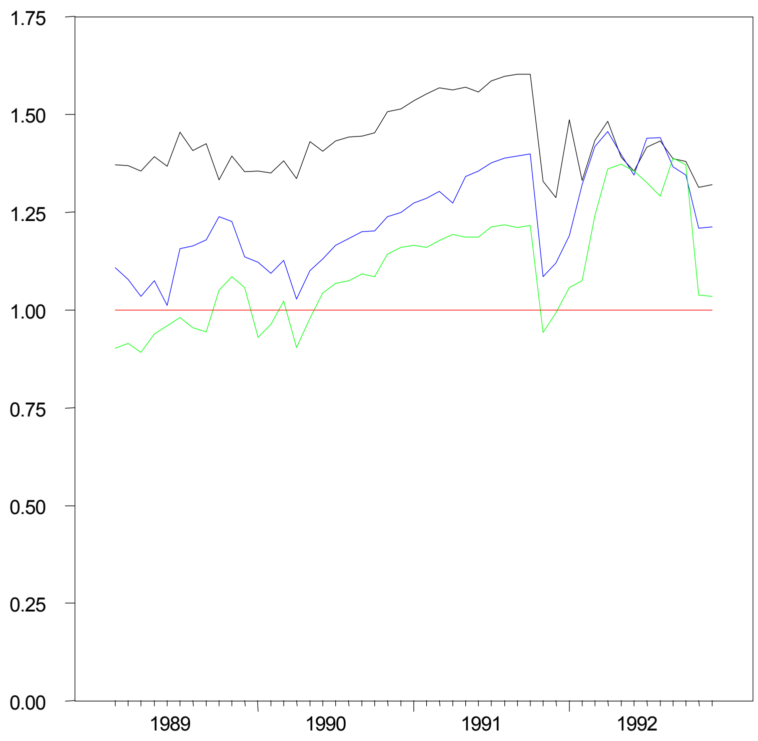


Note: The estimated model tests for a shift in the constant. No trend is included. Sequential statistic are computed using the sample 1984:7-2002:12, sequentially incrementing the date of the hypothetical shift. A fraction equal to 15 per cent of the total sample at the beginning and at the end of the sample is not considered for the test. For more details see Banerjee et al. (1992).

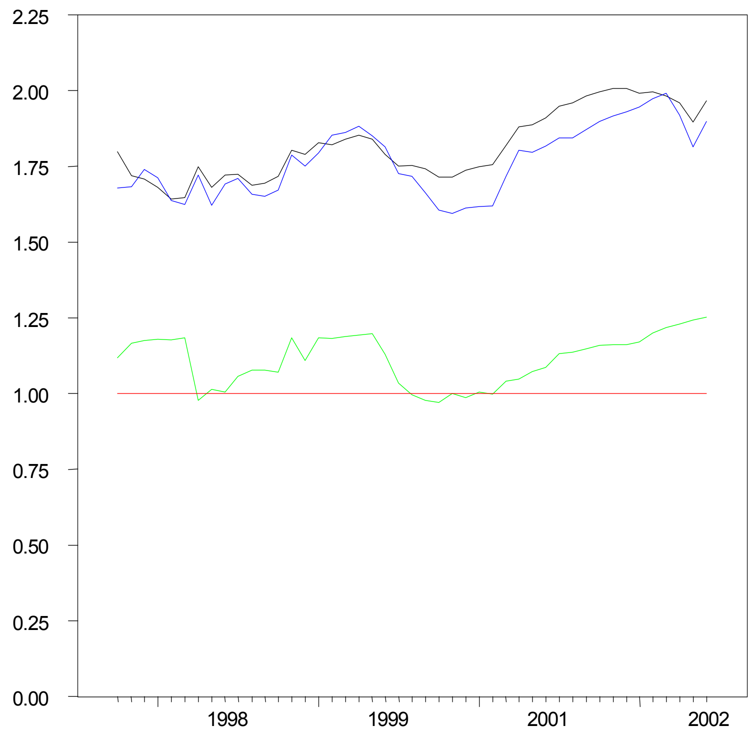
Figure 3

Hansen and Johansen's iterative procedure

I: Partial liberalization period



II: Deregulation period



Note: The above figures represent the cointegrating rank statistics in the two sub-periods (Z-model). They are scaled by their respective critical values (in this way the 95 per cent confidence interval is normalised to one), recursively calculated in the last 4 years of the sample period. Therefore, the lines above the value of one give the outcome of the rank test evaluated for any partial sample period. Due to the presence of dummy and exogenous variables the critical value is only indicative.

Figure 4

Changes in the monetary policy indicator
(percentage values)

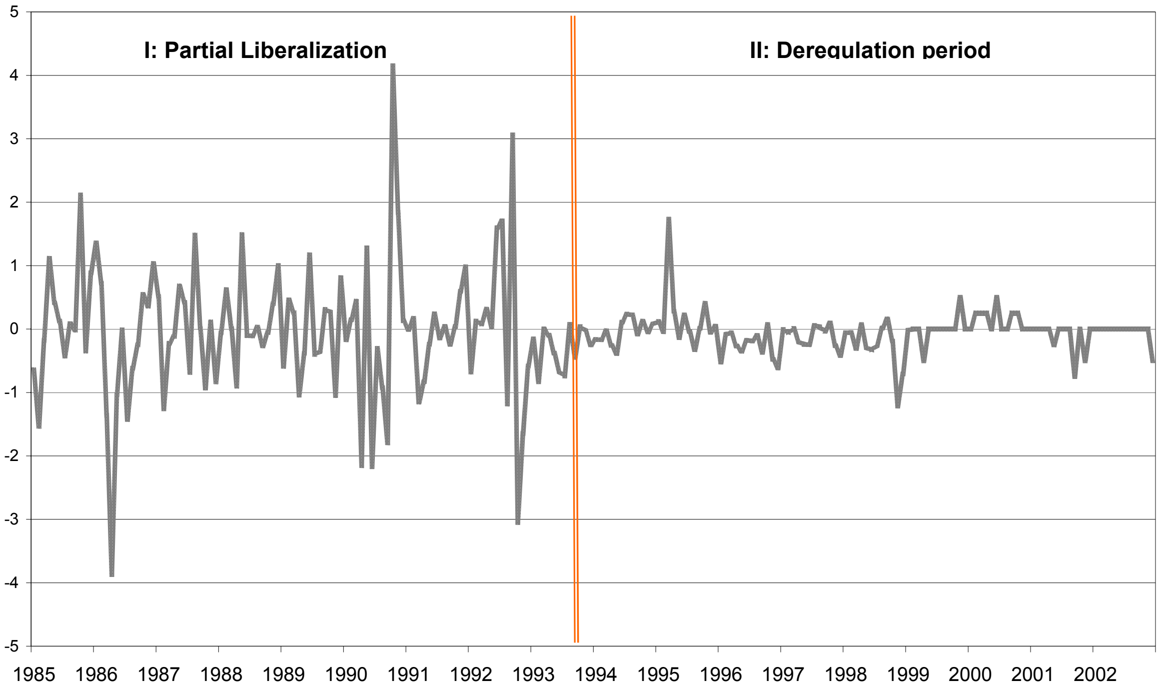
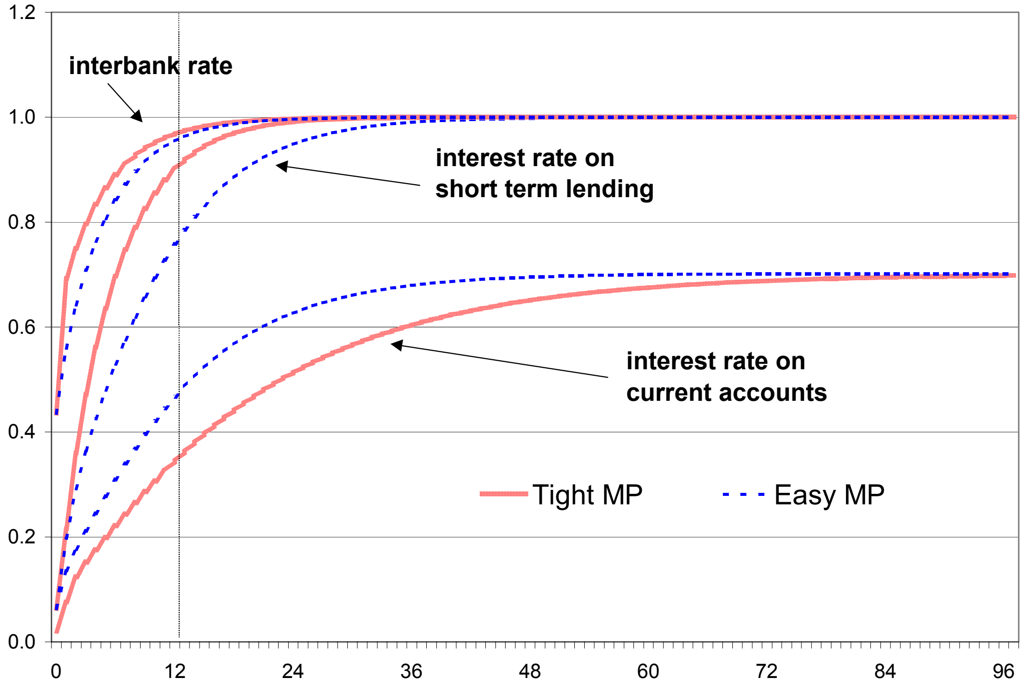


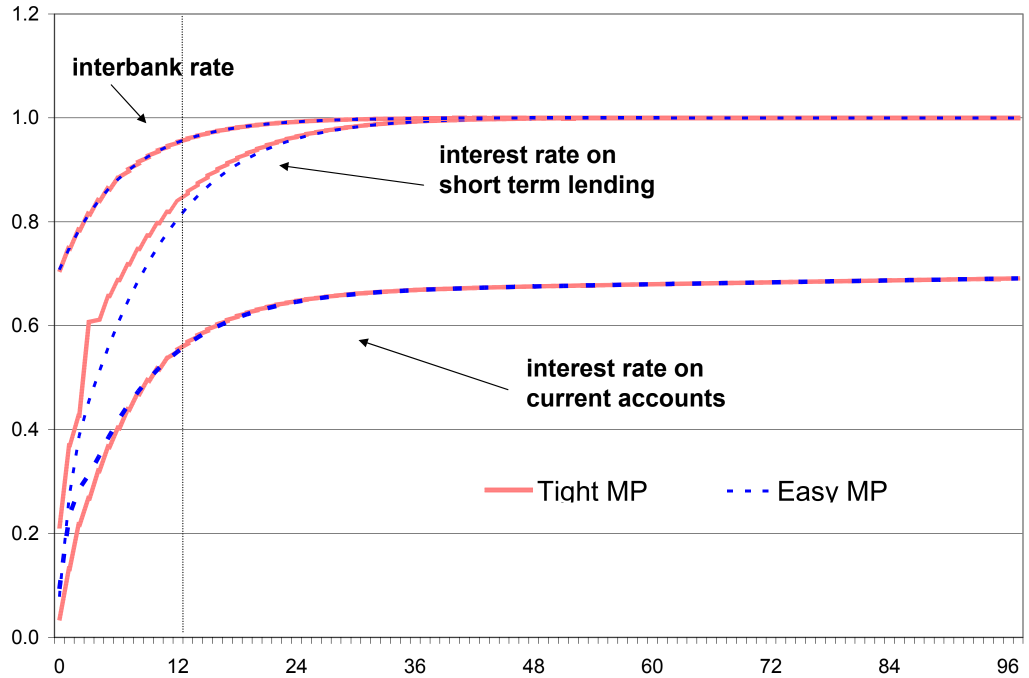
Figure 5

Adjustment paths of bank interest rates to positive and negative changes in the monetary policy indicator
(percentage values)

I: Partial liberalization period (1985:01-1993:08)



II: Deregulation period (1993:09-2002:12)



Note: The experiment consists in evaluating the effect after 96 months of a 1 per cent change in the monetary policy indicator. To make simulations more comparable, the effects for easing monetary policy are multiplied by -1.

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