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Abstract: The empirical literature on interest rate transmission presents diverse and sometimes conflicting estimates. By discussing methodological and specification-related issues, the results of this paper contribute to the understanding of these differences. Eleven Austrian bank lending and deposit rates are utilized to illustrate the pass-through of impulses from monetary policy and banks' cost of funds. Results from vector autoregressions suggest that the long-run pass-through is higher for movements in the bond market than of changes in money market rates. Deposit rates have no predictive content for lending rates beyond that of market interest rates.

Keywords: Monetary policy transmission, interest rate pass-through, retail interest rates, vector autoregression, impulse-response functions.

JEL classification: E43, E52, G21.

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1 Introduction

Numerous studies have provided estimates on the pass-through of interest rates in recent years.[1] In general, retail interest rates (rates that banks charge on loans and pay on deposits from their customers) are found to adjust with a delay to changes in policy (administered or official) and money market interest rates. Completeness of the pass-through in the long run (whether an impulse in the specified reference rate is fully transmitted to retail rates over time), on the other hand, cannot be uniformly confirmed.[2]

The aims of this analysis are twofold. First, as is not generally done thoroughly in similar empirical studies, it shows how pass-through results can be driven by modeling and specification choices. The examined issues also include the importance of in depth data analysis, the choice of proper reference rates and measures of the banks' cost of funds, including a discussion of whether deposit rates have predictive content for lending rates. We additionally deal with the robustness of results from commonly applied bivariate models in multivariate regressions. Based on these results, a better understanding of the (diverse) pass-through estimates to be found in the literature will be possible.

Second, there are some results for Austria, but only for selected lending and deposit rates. As these results are not at all uniform (see Table 1 for a comparison for pass-through results related to the short-term lending rate to enterprises), we seek to assess their reliability, and will provide pass-through estimates for Austrian retail interest rates which were not analyzed before.

The paper proceeds as follows. Section 2 covers some data- and sample-related matters. The handling of suspected structural breaks and the use of single-equation methods are discussed in section 3. Pass-through estimates from bivariate vector autoregressions are presented in section 4. Results obtained in this manner point to an incomplete transmission of interest rate impulses. Adjustment paths for some retail rates are divergent in response to rising or falling reference rates. However, due also to the little number of observations, asymmetry is not confirmed by statistically significant response gaps. Section 5 shows that retail rates respond more strongly to impulses from the bond yield than to shocks from the money market, a result found to be robust in multivariate models. Deposit rates, on the other hand, have no additional predictive content for lending rates when market interest rates are controlled for. Section 6 summarizes and also contains a survey-like assessment of our results and related ones from the literature.

2 Data and related issues

Data on Austrian retail interest rates come from the national interest rate statistics (surveyed before the transition to the harmonized euro area statistics in 2003) and cover the period from March 1995 until June 2003. Seven lending and four deposit rates will be analyzed. Euro area money market rates are appended to the respective Austrian series. For details, also on the bond yield used in the following analysis, see the data description in appendix C.

In choosing appropriate reference interest rates (whose changes are passed through to retail rates), we follow a distinction between impulses in monetary policy and in cost of funds, which also was made in the (euro area) studies of de Bondt (2005) and Sander and Kleimeier (2004a). Overnight money market rates are mainly used as a proxy for the monetary policy stance, since official rates do not change every month. As, for example, the EONIA closely fluctuates around the ECB main refinancing rate (de Bondt 2005), this choice seems warrantable. The above-mentioned studies chose the banks' cost of funds to be represented by the money market rate (MMR) which exhibits the highest correlation with the respective retail interest rate.[3] Sander and Kleimeier (2004a), however, also admit that a market rate with a matched maturity might represent a more favorable explanatory variable in retail interest rate regressions. At the outset, we also select via correlation, but return to the issue in sections 5.2 and 5.3.2.

Interrelations between interest rates only are examined. Therefore, we implicitly assume other factors also affecting retail interest rates to be uninfluential over the sample period. For instance, the degree of competition in the banking sector is exogenous. As shown in Figure 1, there is a downward trend in rates on commercial loans prior to 2000, which was partly driven by increasing competition. Restricting the analysis to a shorter period therefore seems justified (see also section 3.1). During the actual investigated sample, there was one large bank merger (BAWAG and PSK in 2000) which hopefully does not largely affect the results. A second requirement is that the data cover an almost complete business or interest rate cycle which seems to be fulfilled for the period from 1999 to 2003.

Unit root tests (ADF, Phillips-Peron, DFGLS, KPSS and Zivot-Andrews tests) largely indicate that the interest rate series are integrated of order one.

3 Structural breaks and single-equation results

3.1 The breakpoint discussion

The first (still non-methodological) issue to be discussed is the one of structural breaks in the relation between policy (market) and retail rates. A natural candidate with euro area data would be the beginning of 1999. Angeloni and Ehrmann (2003), for example, based on rolling window regressions, conclude that a break occured with the implementation of stage III of the European Monetary Union (EMU). By using Chow tests, de Bondt, Mojon and Valla (2003) find pre- and post-1999 differences in their transmission regressions in about 40% of cases (in 3 out of 4 cases for Austria). On the other hand, Sander and Kleimeier (2004a), when searching for breaks in cointegrating relations, find that these mostly took place before 1999. For Austrian rates on short-term loans to enterprises, for example, the breakpoint found is August 1997. From the first scatterplot in Figure 2 we see that there might have been two breaks in the relation between the commercial lending rate and the overnight MMR. The time period with a negative correlation between the two interest rates starts at the beginning of 1997 and ends in November 1998. So both of the above-mentioned statements about the timing of the break seem to be correct in the Austrian case. As a result of these considerations, we restrict our analysis to the "euro years" (1999-2003).

It seems noticeable that the correlation between the commercial lending rate and the overnight MMR decreases after 1998.[4] The slopes of the corresponding simple regression lines would be approximately 1 (1995-1998) and 0.6 (1999-2003), respectively. A similar picture (which is least pronounced for discount loans, overdrafts and sight deposits) emerges for all lending and deposit rates, independent of which money market rate is on the abscissa.[5] The reasons for these differences in the pass-through seem complex. They might stem from factors driving the downward trend of retail rates in the 1990s (changes in competition, as already mentioned, reductions in operating costs or an increased separation of rates and fees) as well as relating to the EMU. For example, the size of the pass-through is negatively related to interest rate volatility (Sander and Kleimeier 2004a and Mojon 2000), and, as Angeloni and Ehrmann (2003) argue for Germany, the euro area money market rates are more volatile than the Austrian ones were before 1999. However, we will not further investigate the topic in this paper.

The example of commercial credit rate reactions to changes in the overnight MMR also shows that the pass-through estimates are not dramatically influenced by the sample period selection (results not reported). An exception are the cointegration coefficients obtained via the Johansen procedure, which increase to about 1 if the full sample is used. Nevertheless, one should be cautious in comparing results from (euro area) studies which take account for the 1999-break (e.g. de Bondt 2005 or Bredin, Fitzpatrick and O'Reilly 2002) with those of, for instance, Heinemann and Schüler (2002) or Donnay and Degryse (2001), who do not.

3.2 Single-equation models

If time series data are employed, most studies use some form of a single-equation error-correction model (what can also be seen from the results for Austria summarized in Table 1). Sometimes, a vector error-correction model (VECM) is estimated before to obtain the cointegrating relations by the Johansen procedure. Heinemann and Schüler (2002), for example, apply the autoregressive distributed lag (ADL) model[6]

$$R_t = \mu + \beta_0 M_t + \sum_{i=1}^{k} (\alpha_i R_{t-i} + \beta_i M_{t-i}) + \epsilon_t, \tag{1}$$

where R_t stands for the retail and M_t for the policy or money market rate. The impact multiplier is β_0 and the long-run link can be calculated as

$$\theta = \frac{\sum_{j=0}^{k} \beta_j}{1 - \sum_{i=1}^{k} \alpha_i}. \tag{2}$$

Thus, in the long run, the relation becomes

$$R_t = \rho + \theta M_t + u_t. \tag{3}$$

If cointegration between the retail and the reference rate is found to be present, equation (1) is estimated in error-correction form. Then, the long-run pass-through can be directly obtained from equation (3). In case of no cointegration it has to be calculated according to equation (2) or from a corresponding model in differences.

In this paper, we employ vector autoregressive (VAR) models as used also in de Bondt (2005) and Angeloni and Ehrmann (2003) in the context of interest rate transmission. Our estimates of impact multipliers come from contemporaneous (within-month) responses of retail rates to impulses in reference rates. As long-run multipliers we report cointegration parameters and responses to shocks after 12 months.[7]

Additional calculations using ADL models (results not reported) show that both parameters of interest (impact und long-run multipliers) are relatively independent of the lag order choice and of whether we estimate the ADL model in levels or error-correction form. An important point here seems to be that, if equation (1) is estimated in levels, a linear time trend must be included. Multipliers then are similar to those that we obtain from a correctly specified bivariate VAR model (which is discussed in section 4.1). Otherwise, if no trend is included, estimates, especially of long-run multipliers, fluctuate wildly across different possible lag lengths of the model and in some cases are even negative when the lag order is chosen by the Akaike or Schwarz information criterion. Our reasons not to use a single-equation ADL model are stated in section 4.2, and those for not reporting results from equation systems in error-correction form in section 4.1.

4 Retail and money market rates

4.1 Results from bivariate vector autoregressions

Following the recommendations of Ashley and Verbrugge (2004), the VAR models used to calculate impact and long-run multipliers contain both retail and reference rates in levels and a linear time trend. Lag orders are chosen via the Schwarz information criterion. In obtaining structural responses of retail rates to impulses in reference rates, the underidentification problem is solved by the assumption that the reference rate is contemporaneously exogenous.[8] The standard Johansen procedure with unrestricted constants is used to estimate cointegrating relations.

Results for the so-called monetary policy approach (with the monetary policy stance being proxied by the overnight MMR) can be found in Table 2. Simulated retail rate responses always refer to a unit (one percentage point) shock in the overnight interbank rate. Cointegration parameters and our other measure for the long-run multiplier, the response of the respective retail rate after 12 months, are of similar magnitude for all lending and deposit rates. For the five main lending rates, the impact multiplier is around 0.4 and the long-run multiplier around 0.6. The interest rate on discount loans shows a stronger adjustment (around 0.8 in the long run) and also the lower multipliers for overdrafts, sight and overnight deposits were as expected. Interest rates on longer-term savings adjust by about 0.3 percentage points within the month the shock occurs, and by 0.4 in the long run. Figure 3 shows, as one example, the response of the commercial loans rate to a unit shock in the overnight MMR, for the month the shock occurs as well as for the twelve following months.[9]

Banks' cost of funds are, as in de Bondt (2005) and Sander and Kleimeier (2004a), represented by the interbank rate which features the highest correlation with the respective retail rate. The estimated impact and long-run multipliers can be found in Table 3.[10]

Cointegration parameters for the five main lending rates with the respective reference interest rate are again around 0.6. Some of the impact multipliers after a unit shock are lower, some are higher than the comparable responses after an impulse in the overnight market rate. Long-run multipliers for the rates on commercial and mortgage loans appear to be somewhat higher than with the monetary policy approach.

Some specification changes to explore the robustness of these results are insightful. The first example is the effect of different lag lengths, illustrated here by dint of the commercial lending rate (monetary policy approach). Cointegrating relations are rather unperturbed by way of changing VAR orders, but the impulse response functions are affected in two ways. The responses to shocks in the reference rate, in general, are higher with more lags, but with more lags they are likely to exhibit a hump shape in time. The maximum response, for example, of the commercial credit rate to shocks in the overnight MMR is near 1 if three lags of each variable enter the VAR. The AIC would choose 2 lags, the maximum response then is 0.8 and the long-run response would amount to 0.7.

Second, there shall be some comparison to results from vector error-correction models (VECM), which are, for most of the cases in this paper, not equivalent to the results from VARs. Two factors are responsible for these differences. If, on the one hand, the error-correction model should include one lag, the corresponding level-VAR has an order of two. The Schwarz information criterion, however, mostly leads us to set the VAR order to one. On the other hand, error-correction models may suffer from inappropriate cointegrating relations which lead to biased response functions.

The relation between the 1-month MMR and the rate on mortgage loans provides the illustrative example because the Schwarz criterion suggests the level-VAR order to be two in this case. The accumulated response after 12 months from the VECM is 1.46 (which is rather high when compared to the mortgage loans rate response of 0.85 from Table 3). The sum of the significant responses is 1.09 since only the first five responses after the shock are statistically different from zero (at the 5% level). We conclude (also for the remaining retail rates) that vector error-correction models do not perform satisfactorily in obtaining reliable transmission estimates.

4.2 Results on asymmetry of responses (in bivariate models)

The evaluation of response asymmetry is an integral part of the agenda in studying the interest rate pass-through. Whether retail rates respond differently to rising and falling reference rates, or when (changes in) banks' cost of funds exceed a certain threshold, may be analyzed in various ways. Several authors differentiate between sample periods of only rising or falling policy (market) rates (for example Mojon 2000, Heinemann and Schüler 2002 as well as Mester and Saunders 1995) or use interaction terms (Rosen 2002). In most cases, however, the assessment of response asymmetry is based on the error-correction form of equation (1), that is

$$\Delta R_t = \mu + \beta_0 \Delta M_t + \sum_{i=1}^{k-1} (\alpha_i \Delta R_{t-i} + \beta_i \Delta M_{t-i}) + \gamma ECT_{t-1} + \epsilon_t, \tag{4}$$

where $ECT_t = R_t - \rho - \theta M_t$ represents the deviations from the long-run relationship from equation (3). The simplest asymmetric specification then is

$$\Delta R_t = \mu + \beta_0 \Delta M_t + \sum_{i=1}^{k-1} (\alpha_i \Delta R_{t-i} + \beta_i \Delta M_{t-i}) + \gamma_1 ECT^{+}_{t-1} + \gamma_2 ECT^{-}_{t-1} + \epsilon_t. \tag{5}$$

The error-correction term is split, so that $ECT^{+}_{t} = ECT_t$ if there is a positive equilibrium error, and zero otherwise. Similarly, $ECT^{-}_{t} = ECT_t$ for negative deviations from the cointegrating relation.[11]

Asymmetry of retail rate adjustment is only allowed for deviations from the long-run relation. However, the more serious problem with the asymmetric ECM seems to stem from the accepted interpretation of the adjustment coefficients. Positive equilibrium errors, in general, are interpreted as representing times of falling policy or market rates. Negative deviations are proposed to stem solely from a more restrictive monetary policy or cost increases in the interbank market.

Figure 4 illustrates an example of potentially associated misinterpretations. The first graph shows that the deviations from the estimated cointegrating relation between the overnight MMR and the short-term commercial lending rate (from 1999 to 2003) are positive, above all, at times when the interbank rate rises, and below zero at times of a falling market rate. This is diametrically opposed to the above intention in interpreting equation (5). Figure 5 shows a scatterplot of the equilibrium error against the change in the overnight MMR, which suggests a positive correlation between these two measures. A negative one would suit the hypotheses discussed here.[12]

As our approach is to estimate a VAR in levels of the variables, we have to apply a different method to investigate asymmetry anyway. In Tables 3 to 5 (column headers rising and falling MMR) we quote retail rate responses from a VAR where the retail interest rate equation is specified as

$$R_t = \mu + \sum_{i=1}^{k}(\alpha_i R_{t-i} + \beta_i M_{t-i} + \gamma_i M_{t-i} I_{t-i} + \delta_i I_{t-i}) + \phi t + \psi Z + \epsilon_t \tag{6}$$

and the identification scheme again assumes that the MMR contemporaneously affects the retail rate but not vice versa. The indicator function (dummy) I is defined to represent cases of rising (falling) interbank rates so that the β coefficients measure the effects of falling (rising) rates.[13] Z is for additional deterministic terms (dummies).[14] Similar specifications are used by Hofmann and Mizen (2004) and Tkacz (2001) in error-correction models as well as by Lim (2001) for a vector error-correction model.

According to, for example, Hannan and Berger (1991), lending (deposit) rates should be more sticky when the interest rate level goes down (up), except for when the threat of negative customer reactions is significant. They, however, do not find evidence on the role of the latter as a major cost of changing prices. A comprehensive assessment of the literature regarding the results on response asymmetry and their statistical and economic significance would go beyond the scope of this section. Below, comparisons will be limited to relating our results to those of other papers providing estimates for Austrian retail rates.

A look at the asymmetry results in Tables 2 and 3 (long-run multipliers only[15]) reveals that asymmetry is largely an issue with lending rates. After impulses in the overnight MMR (Table 2), responses are higher if the reference rate goes up than for decreases. However, the respective long-run multipliers do not seem to be significantly different as the confidence bands (see note 9) overlap. For illustration purposes, Figure 6 shows the responses of the consumer loans rate (for which the estimated gap is the widest). Even in this case, we should therefore be cautious of arguments in favor of asymmetry.

The interest rate on commercial loans seems to adjust more strongly after decreases in the reference cost-of-funds rate. This result in Table 3 seems to be driven by three periods (end of 1999, mid of 2000, beginning of 2002) where the 3-month MMR rose and the commercial lending rate increased only very little or even continued to fall for some subsequent months. So we do not suppose that this reversed asymmetry results from the threat of possible negative consumer reactions that may force banks to adjust their lending rates more and more quickly if rates in general go down.

Similar responses also appear within the results for Austrian (mainly lending) rates in de Bondt et al. (2003). They apply the splitting technique with respect to the error-correction term (see equation (5)) and often find that the adjustment back to equilibrium is faster in case of positive deviations from long-run relations. So their results seem to suffer from the above-mentioned fact that positive equilibrium errors need not be associated with times of falling market rates. Surprisingly, the Austria-results of Sander and Kleimeier (2004b) (this working paper provides the most details about their studies) are not affected likewise (there are some suspicious table entries for other countries). Although they do not offer standard errors for their estimates, the differences between responses after simulated positive and negative 1% shocks to the respective money market rate do not seem to be significantly different. So, within the scope of bivariate models, it seems that our results coincide with the ones Sander and Kleimeier (2004b) report for Austrian data.

5 Bond yields and results from multivariate models

This section covers some issues which are not generally discussed in each article on interest rate transmission. Also the robustness of our previous results to some changes in the model specification is considered. For the most part, we leave the bivariate models commonly applied.

5.1 Responses to cost-of-funds changes if policy is controlled for

For the five main lending rates, we calculated responses to changes in the cost of funds that are not correlated with shocks in the overnight MMR. Market rates representing these costs are, as before, chosen via their correlation to the respective lending rate. The applied identification scheme for impulse-response analysis assumes that, contemporaneously, the overnight market rate is exogenous and has an influence on both the cost-of-funds rate and the retail rate. The longer-term MMR only affects the lending rate (but not vice versa). So Table 4 reports the lending rates' responses to changes in the banks' cost of funds which are uncorrelated with (orthogonal to) shocks in monetary policy. These turn out to be hump-shaped in time in many cases, so the maximum responses during the 12-month-period are quoted.

Responses of lending rates are now estimated to be higher, not very different from one. The highest long-run adjustment is the case for interest rates on commercial and consumer loans. A similar model (in error-correction form) is estimated by de Bondt (2005) for the euro area aggregate lending rates, yielding fairly comparable results. Lending rates respond fully in the long run to changes in market rates, if the model is specified as such. The same is true for the retail rate adjustment estimates of Borio and Fritz (1995) and Cottarelli and Kourelis (1994). For most of the countries they consider, the long-run pass-through to short-term firm lending rates is approximately one.

5.2 Maturity matters

As de Bondt (2005) argues, the reason for the pass-through to long-term bank lending rates often being found to be less complete than for short-term bank lending rates (to enterprises) might be "that the marginal cost prices are approximated by money market interest rates which may not be the most appropriate marginal funding costs for long-term loans".

We presume the yield of bonds issued by banks[16] to be an adoptable proxy for banks' cost of funds or the "maturity structure of bank balance sheets" (de Bondt et al. 2003). Results from bivariate vector autoregressions are reported in Table 5. The null of no cointegration is always rejected (with very high values of the trace statistic) and almost all long-run coefficients are around one. Retail rate adjustment after 12 months, in general, is lower than one, but around 0.8 for lending rates. So we confirm that "maturity matters" (de Bondt 2003) by finding that retail rates respond more to changes in bond yields than to movements in money market rates. Section 5.3.2 will show the robustness of this result.

5.3 Results from multivariate vector autoregressions

5.3.1 Are deposit rates Granger causal for lending rates?

Next, we examine multivariate models of interest rate transmission. For the five main lending rates, we here first pose the question of whether a multivariate description of lending rates must also consider deposit rates.[17]

Studies measuring banking activity, productivity and efficiency are related to this issue. By that literature, the question of whether deposits are inputs or outputs in the production process of a banking firm has been discussed extensively. The so-called asset or intermediation approach argues that deposits are inputs to loan-making. If, on the other hand, the role of deposits as a service to the banks' customers is emphasized (so deposits are seen as outputs), we speak of the so-called production or service provision approach (Mlima and Hjalmarsson 2002). Data-driven methods to evaluate these views mostly conclude that deposits should be classified as outputs (Fixler and Zieschang 1999).

As de Bondt, Mojon and Valla (2002) argue, the rates on loans granted by universal banks (opposed to those of specialized banks) "may depend on the cost of raising deposits rather than issuing securities. Such a deposit-based funding of loan activities could imply that retail bank rates remain little responsive to market conditions once deposit rates are accounted for. On the contrary, specialised banks without branches collecting deposits would set their retail loan rates on the basis of their market-based funding."

As we presume the Austrian banks surveyed for our interest rate data to be such universal banks[18] we seek to explore how deposit rates are related to lending rates. It turns out that, among the deposit rates, the one for savings deposits with an agreed maturity of over 12 months has the highest correlation with each of the five major lending rates. We therefore test whether this savings rate has predictive content for lending rates beyond the policy and cost-of-funds measures employed so far.

Four variables (lending, policy, interbank and deposit interest rates) enter the VAR being the framework for Granger-noncausality tests. We follow Toda and Yamamoto (1995) in adding one augmenting lag to the dynamic structure, which is not used by the test but enables valid Granger-noncausality inference to be conducted in models that contain unit roots.

Our results (see Table 6) do not support Granger causality of deposit rates for interest rates on loans to enterprises as well as consumer credit, a finding of de Bondt et al. (2002, 2003). For rates on housing, mortgage and municipal loans there is also no such effect. Consequently, we find that all respective lending rate responses to deposit rate changes are insignificant for the months after the shock (results not reported). Contemporaneous responses, however, are significantly different from zero for consumer und municipal credit.[19]

Ultimately, the results here tend to support that deposits are to be classified as outputs of banks' production process. We therefore do not account for interest rates on savings when describing lending rates in the remainder of the analysis.

5.3.2 Cost-of-funds considerations in multivariate models

The last step in our assessment of the interrelations of retail to policy and other interest rates that proxy the banks' cost of funds is a multivariate VAR including a retail rate, the overnight and the selected longer-term MMR, and the bond yield. Some results of Granger-noncausality tests (as described in section 5.3.1) can be found in the second part of Table 6. The overnight and the longer-term MMR are not Granger-causal for lending and saving rates, except for the lags of the 1-month MMR having some predictive content for housing and mortgage loan rates (the respective test results are not reported).

However, the bond yield can be said to Granger-cause all retail rates in this multivariate setting. Another result is that the long-run responses of retail rates to impulses in the bond yield from Table 5 are relatively robust to the inclusion of other variables.[20] In the multivariate setting, the long-run adjustment of rates on mortgage and municipal loans to changes in the bond yield are somewhat higher than before, those of the longer-term savings rates are somewhat lower. For a graphical example see Figure 7, which contains the responses of the interest rate on mortgage loans to a unit shock in the bond yield.

From our results it seems evident that the bond market is more important for transmitting interest rate impulses than the money market, suggesting that the pass-through process might have a two-stage nature (similar to the one described by de Bondt 2005). Short-term money market rates first affect market rates with longer maturity, and then retail rates respond to those changes in longer-term market rates. If this is a sound description of reality, there should be no additional long-run effect of short-term interbank rates in models that involve bond rates.[21] As this is the case, we support Sander and Kleimeier (2002) in that it appears that such an announcement effect (as described by Cottarelli and Kourelis 1994) "has lost in importance as banks increasingly seem to base their pricing decision on cost of funds considerations."

6 Summary, comments and conclusions

This study has examined various issues raised in the empirical literature on the transmission of interest rate signals to banks' retail rates. Thereby, we followed de Bondt (2005) and Sander and Kleimeier (2004a) in their distinction of a monetary policy and a cost-of-funds approach. Simple data analysis led us to account for more than one structural break in interest rate relations and to restrict our sample period to the euro years. The inclusion of a linear time trend seemed warranted in single-equation as well as simultaneous-equation models. Different choices of the lag order were also found to produce significantly different results. Section 4 also illustrated why VARs are to be preferred to VEC models and that using deviations from cointegrating relations in asymmetry analysis is often questionable. In light of these and some points to be discussed below, we derived a new set of results for Austria which is more reliable than previous ones. An additional contribution of this paper is that it provides interest rate pass-through estimates for a total of eleven retail rates.

The results from bivariate models suggest severe price rigidities in the banking sector after changes in marginal funding costs proxied by movements of rates in the money market. As Sander and Kleimeier (2004a) argue, country-specific banking-market and legal characteristics make for Germany and Austria typically having a lower transmission of interest rate impulses than other countries. In the end, we can confirm the results of Sander and Kleimeier (2004a) who showed that (for Austria, in bivariate models with interbank rates) the long-run pass-through for most of the retail rates is found to be incomplete. Even many coefficient magnitudes are similar despite several methodological differences. Sander and Kleimeier (2004a) argue that this result "... does not depend on the choice of the market rate proxy and is thus standing in contrast to the studies by de Bondt (2002) and de Bondt et al. (2002)."

But whether these results really are at odds remains unclear. First, a complete pass-through only is the case for lending rates in de Bondt et al. (2002) and de Bondt (2005), not for deposit rates. Second, de Bondt (2002, 2005) examines impulse responses from vector autoregressive models calculated for 36 months following the shock. Responses of lending rates after 12 months are also (partly much) lower than one, of similar magnitude to our estimates for Austria. We suppose that the transmission process is almost completed after one year for two reasons. Indeed, retail rate responses still increase afterwards, but the respective incremental changes are small and typically estimated very imprecisely. Additionally, the adjustment progress after 12 months is pretty much in line with the estimated cointegration (long-run) parameters.

As we saw in section 5.1, another reason for studies reporting that "there is a close to one-to-one pass-through in the long term of changes in market interest rates to retail bank interest rates" (de Bondt 2003), is that of controlling for a proxy of monetary policy. But also in bivariate models we observe estimates of long-run transmission of about one. Examples are Toolsema, Sturm and Haan (2002), Sander and Kleimeier (2002), Angeloni and Ehrmann (2003) and Hofmann and Mizen (2004). As we have seen that several modeling and specification choices (treatment of structural breaks, choice of the lag order, etc.) can lead to rather large changes in estimates, their results need not, from the outset, contradict those of de Bondt (2005), Sander and Kleimeier (2004a) and this paper.

On the other hand, the evidence presented here leads us to conclude that we should take a look beyond bivariate models with retail and money market rates anyway. First, impulses in money market rates that are not correlated with monetary policy shocks are fully transmitted to retail rates. Second, bond yields are expected to provide a better maturity matching to most lending and deposit rates. Therefore, they should describe cost-of-funds changes for banks in a more adequate way than interbank rates. Granger causality results from multivariate models confirm that money market interest rates do not have predictive content for retail rates once bond yields are controlled for. Pass-through estimates from bivariate and multivariate specifications suggest an almost complete long-run transmission of shocks in the bond market, at least to lending rates.

Notes

[1]See, for example, Hofmann and Mizen (2004) for a general discussion of the pass-through issue.

[2]Sander and Kleimeier (2004a) argue that differences in pass-through estimates across the literature can mainly be ascribed to the chosen methodology, the choice of the exogenous market rate, the length and timing of the sample period and the treatment of possible structural breaks. Regarding the latter points, a question of interest is often whether transmission has become quicker since 1999 (start of stage III of the EMU).

[3]Banks are generally presumed to operate under imperfect competition on markets for loans and deposits. Assuming perfect competition in the banking sector is not appropriate because of important barriers to entry (Freixas and Rochet 1997) and product differentiation (Hannan and Berger 1991). So banks have some power in price setting (Hofmann and Mizen 2004) for these products, or are assumed to follow a markup pricing rule (Winker 1999). The reactions to changes in an exogenous measure of banks' cost of funds then point to the existence of a cointegrating relation between market and retail rates (see, for example, Lim 2001).

[4]Modeling and robustness issues are illustrated considering the commercial lending rate as an example. If no contrary statements are made, the results from these exercises can be generalized to the other retail rates.

[5]No such obvious breaks can be observed in relations of lending rates with deposit rates (which is supportive of our results in section 5.3.1) or the bond yield.

[6]Sander and Kleimeier (2002, 2004a) allow for a different lag length for the two variables. Cottarelli and Kourelis (1994) and Toolsema et al. (2002) employ a similar equation with one autoregressive lag only.

[7]One feature of the ADL model is that the reference interest rate (the one the impulse comes from) is a priori treated as exogenous. In the bivariate case, our impulse-response analysis assumes that the reference rate is contemporaneously exogenous. Also Donnay and Degryse (2001) use a VAR, but do not simulate responses to one-time unit shocks.

[8]Expressed in technical terms, this amounts to using the Cholesky decomposition of the variance-covariance matrix of the VAR residuals to identify the structural shocks via the reduced-form VAR estimates. The Cholesky decomposition involves a causal chain (recursiveness) in contemporaneous relations between the variables.

[9]All impulse-response graphs contain the tighter bands from the 0.16 and 0.84 fractiles of the response distribution, although statements about statistical significance in the text refer to the 5% significance level (or, more exactly, the 0.025 and 0.975 fractiles). Adaptations of the Rats example programs monteva2 and montesur (from estima.com) were used to calculate impulse responses and corresponding error bands.

[10]Table 3 (as Table 4) only reports results for the five main lending rates. The discount loan and the longer-term deposit rates show correlations which are decreasing with the maturity of the corresponding interbank rates. Interest rates on overdrawing and the shorter-term deposit rates are left out because they have relatively little variation.

[11]More sophisticated threshold cointegration models (see Balke and Fomby 1997 for EQ- and B-TAR, as well as Enders and Siklos 2001 for TAR and M-TAR) are also employed in the interest rate transmission context (Sander and Kleimeier 2002, 2004a,b, Tkacz 2001). If asymmetries are present, estimating these models may facilitate the detection of cointegration (Sander and Kleimeier 2002). The asymmetric model of equation (5) can be seen as implying a threshold value of zero. Scholnick (1996, 1999) and de Bondt et al. (2003) are examples of its empirical application.

[12] For other retail rates, and when applying the cost-of-funds approach, similar pictures emerge, though they are not as pronounced as in Figure 5 in every case.

[13] Equation (6) actually corresponds to a specification with interactions or, in technical terms, to a threshold model with ΔM as the threshold variable. The threshold value is implicitly zero, and different intercepts and slopes in different regimes are allowed for.

[14] Our asymmetric specification is also not a priori resistant to the interpretation problems described for error-correction terms. Nevertheless, accounting for some obvious data features sometimes provides remedy. For example, there is an obvious shift in the relation between the corporate lending rate and the overnight MMR (see the second scatterplot in Figure 2). We therefore enrich the VAR with a simple shift dummy. Most of the other relations also exhibit such shifts.

[15] Impact multipliers (within-month responses) for the asymmetric models in Tables 2 to 5, on the whole, are of similar magnitude as the ones from symmetric VARs. Generalized impulse-response functions (see Koop, Pesaran and Potter 1996 and Pesaran and Shin 1998), which are to be preferred in nonlinear models, were also calculated. In general, these responses are somewhat smaller than the reported ones.

[16] Results not reported here reveal that transmission estimates from changes in the 12-month MMR are not really different from the ones reported so far (with the 1-month or 3-month MMR proxying the cost of funds).

[17] The importance of deposits for loan funding can be seen if we break down the total financial liabilities of Austrian monetary financial institutions into the different instruments: Currency 2.6%, securities 22.5%, deposits 64.4%, loans 3.4%, shares and other equity 6.2%.

[18] This is supported by the description of the sample in Klein, Schubert and Swoboda (2003). Also de Bondt et al. (2002) suppose most of the European banks to be universal and thus to have a multi-business nature.

[19] Lending rate responses to deposit rate shocks from bivariate vector autoregressions (results not reported) are high contemporarily (0.70 to 0.98) and fall continuously in time (become insignificant after 3-5 months). We consider this as additional evidence of deposits being outputs in bank production.

[20] This is independent of their place in the ordering (here, the ordering is overnight market rate, longer-term interbank rate, bond yield, retail rate). Generalized impulse responses, which are order-invariant, are also in this case not largely different to the reported ones.

[21] The vector error-correction specification of de Bondt (2005) rules out such an effect by restricting the coefficient of the official interest rate to zero in the cointegrating relation, which is consequently formed by market and retail rates.

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A Figures

Figure 1: The interest rate on loans to enterprises and the overnight money market rate.

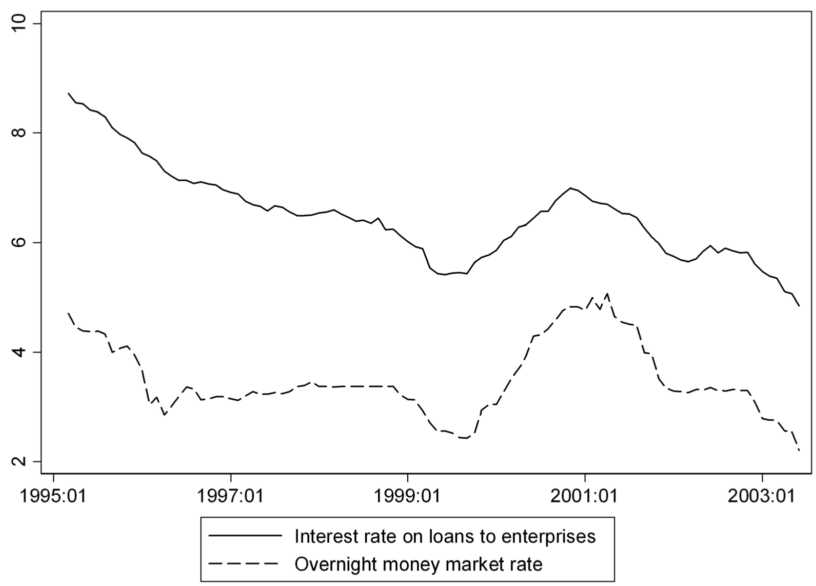


Figure 2: The interest rate on loans to enterprises vs. the overnight money market rate, 1995:03-2003:06 (top), 1999:01-2003:03 (bottom).

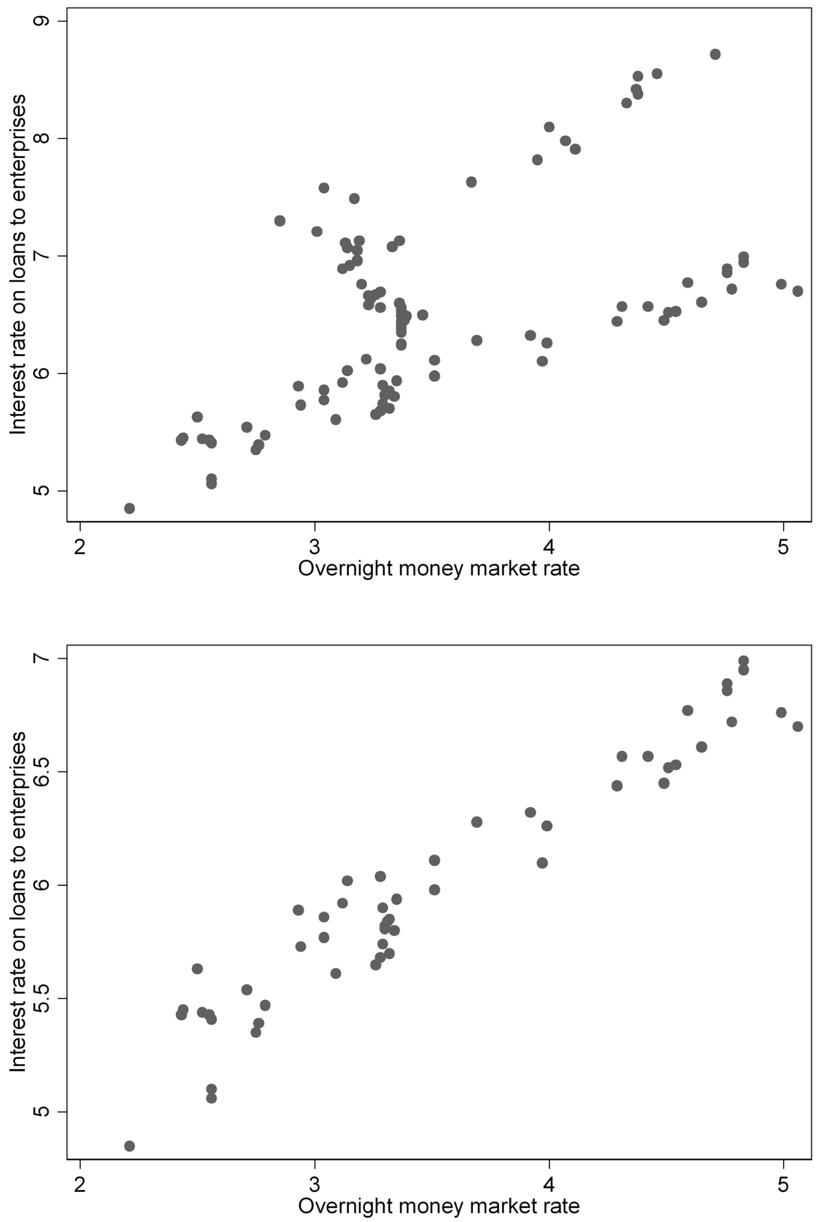


Figure 3: The response of the interest rate on commercial loans to a unit shock in the overnight money market rate.

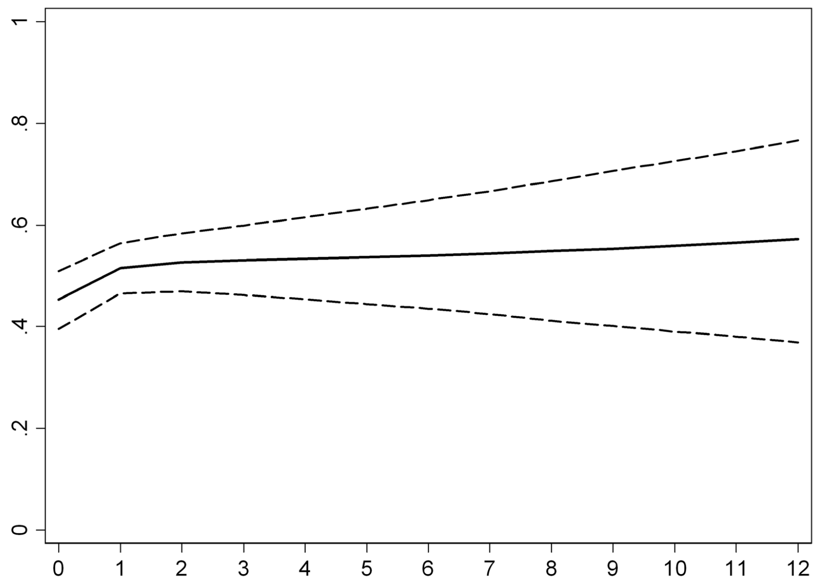


Figure 4: The cointegrating relation between the interest rate on loans to enterprises and the overnight money market rate.

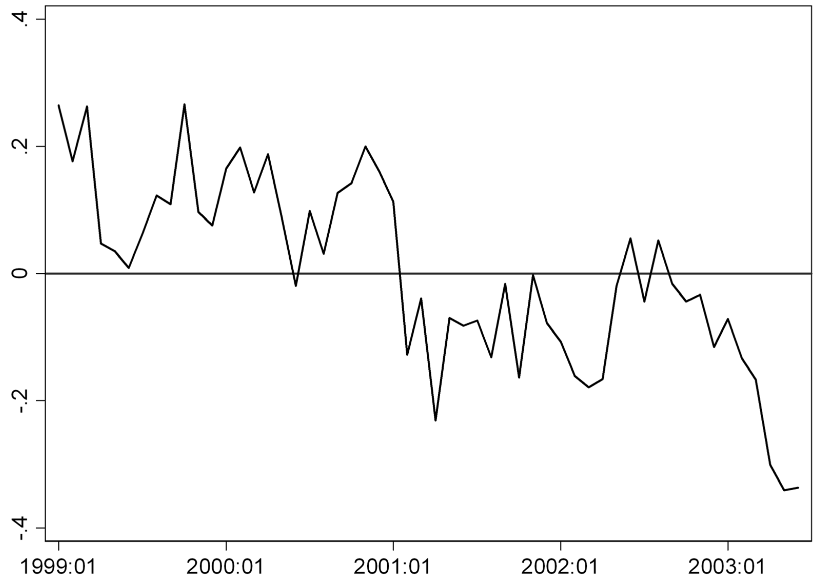


Figure 5: The cointegrating relation between the interest rate on loans to enterprises and the overnight money market rate vs. the change in the overnight MMR.

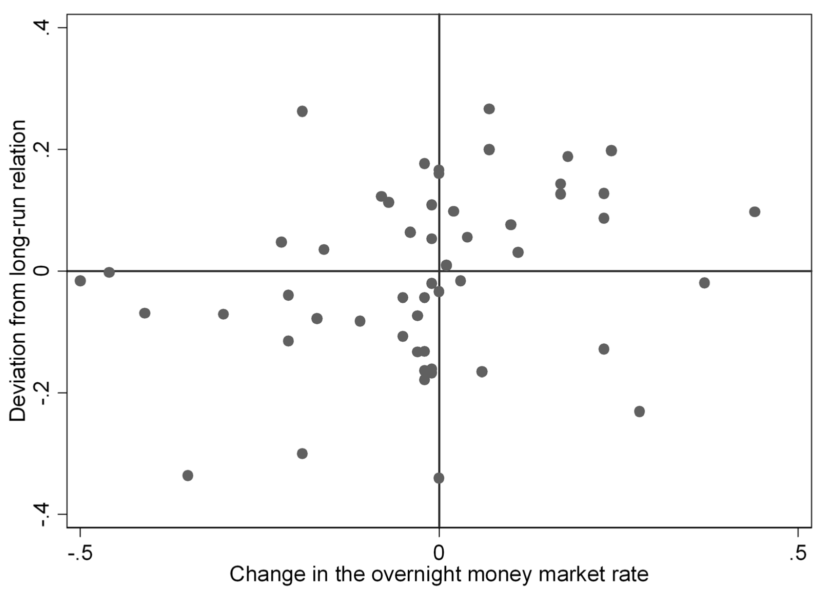


Figure 6: The response of the interest rate on consumer loans to an increase (top) or a decrease (bottom) of the overnight money market rate.

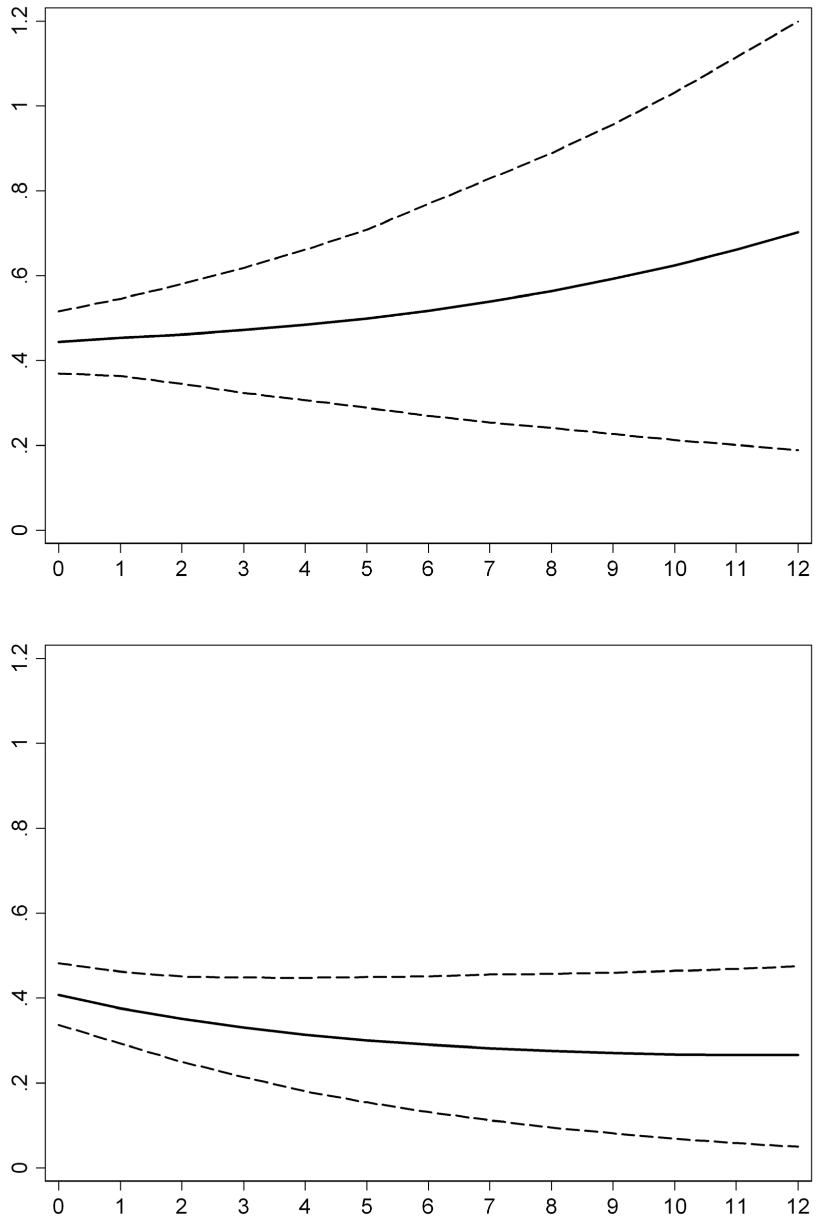
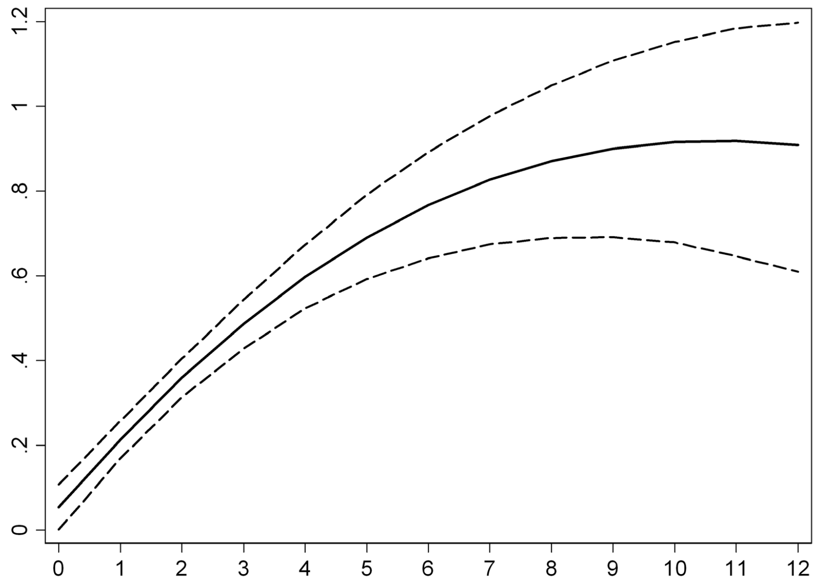


Figure 7: The response of the interest rate on mortgage loans to a unit shock in the bond yield.



B Tables

Table 1: Pass-through results for Austria: Interest rate on short-term loans to enterprises.

Author(s)	Impact multiplier[a]	Long-run multiplier[a]	Method[b]	Sample period	MMR[c]
Donnay and Degryse (2001)	0.08	0.18	VAR	95-00	overnight
Sander and Kleimeier (2002)	-0.18	0.22	ECM	94-98	overnight
de Bondt et al. (2003)	0.24	0.38	ECM	95-02	3-month
	0.38	0.62	ECM	99-02	3-month
Sander and Kleimeier (2004b)	0.03	1.02	ECM	95-97	overnight
	0.24	0.52	ECM	97-02	overnight
	0.05	1.19	ECM	95-97	1-month
	0.26	0.56	ECM	97-02	1-month

[a] Definitions of impact and long-run multipliers differ and, therefore, the results are comparable with reservation only.
[b] The error-correction models (ECM) typically used are symmetric and linear single-equation models containing a contemporaneous term of the money market rate change (the latter is treated as being also contemporaneously exogenous).
[c] The money market rate chosen to proxy either the monetary policy stance or the banks' cost of funds.

Table 2: Monetary policy approach.[a]

Interest rate	Cointegration parameter	Impact multiplier	Long-run multiplier[b]	Rising MMR[c]	Falling MMR[c]
Commercial loans	0.61	0.45	0.57	0.59	0.47
Consumer loans	0.60	0.46	0.58	0.70	0.27
Housing loans	0.59	0.39	0.53	0.56	0.40
Mortgage loans	0.57	0.38	0.58	0.73	0.52
Municipal loans	0.58	0.36	0.62	0.75	0.47
Discount loans	0.79	0.27	0.81	0.89	0.67
Overdrafts	0.16[d]	0.18	0.20	0.21	0.13
Sight deposits	0.05[d]	0.02	0.03	0.03	0.03
Overnight savings	0.12	0.09	0.10	0.09	0.12
Savings ($<$ 12 months)	0.45[e]	0.29	0.44	0.44	0.43
Savings ($>$ 12 months)	0.47	0.26	0.43	0.39	0.47

[a] The overnight money market rate reflects the monetary policy stance. The VAR order is one in each case.
[b] Response after 12 months.
[c] Responses after 12 months from the corresponding threshold model.
[d] Deviations from the cointegrating relation are not found to be stationary.
[e] Deviations from the cointegrating relation are stationary at the 10% level.

Table 3: Cost-of-funds approach.[a]

Interest rate	Cointegration parameter	Impact multiplier	Long-run multiplier[b]	Rising MMR[c]	Falling MMR[c]
Commercial loans	0.65	0.36	0.71	0.66	0.78
Consumer loans	0.61	0.50	0.59	0.61	0.43
Housing loans	0.62	0.34	0.62	0.63	0.61
Mortgage loans	0.60	0.27	0.72	0.73	0.70
Municipal loans	0.62[d]	0.41	0.60	0.59	0.60

[a] Cost of funds are proxied by the 1-month money market rate except for commercial and municipal loans (3-month MMR). The VAR order is one in general, and two for mortgage loans.
[b] Response after 12 months.
[c] Responses after 12 months from the corresponding threshold model.
[d] Deviations from the cointegrating relation are stationary at the 10% level.

Table 4: Cost-of-funds approach II.[a]

Interest rate	Impact multiplier	Long-run multiplier[b]	Rising MMR[c]	Falling MMR[c]
Commercial loans	0.30	1.02	1.03	1.00
Consumer loans	0.52	1.02	1.02	0.90
Housing loans	0.36	0.81	0.84	0.78
Mortgage loans	0.33	0.85	0.85	0.81
Municipal loans	0.41	0.90	0.90	0.85

[a] Costs of funds are proxied by the 1-month money market rate except for commercial and municipal loans (3-month MMR). The VAR order is one.
[b] Maximum response.
[c] Maximum response from the corresponding threshold model.

Table 5: Cost-of-funds approach III.[a]

Interest rate	Cointegration parameter	Impact multiplier	Long-run multiplier[b]
Commercial loans	1.12	0.13	0.78
Consumer loans	1.04	0.07	0.80
Housing loans	1.08	0.07	0.85
Mortgage loans	1.04	0.07	0.84
Municipal loans	1.13	0.24	0.82
Savings ($<$ 12 months)	0.87	0.10	0.57
Savings ($>$ 12 months)	1.02	0.05	0.69

[a] Costs of funds are proxied by the bond yield. The VAR order is one.
[b] Response after 12 months.

Table 6: Multivariate analysis.[a]

Interest rate	Deposit rate: Granger non-causality[b]	Deposit rate: Impact multiplier[c]	Bond yield: Granger non-causality[b]	Bond yield: Long-term multiplier
Commercial loans	0.40	0.25	0.06	0.74
Consumer loans	0.72	0.37 **	0.02	0.83
Housing loans	0.13	0.24	0.00	0.88
Mortgage loans	0.22	0.25	0.00	0.91
Municipal loans	0.76	0.51 **	0.01	1.10
Savings ($<$ 12 months)			0.00	0.44
Savings ($>$ 12 months)			0.00	0.61

[a] The VAR order always is one.

[b] P-value of the F-test on Granger non-causality in a VAR including both the overnight and the longer-term money market rate which shows the highest correlation with the respective lending (deposit) interest rate.

[c] Two asterisks indicate statistical significance at the 5%-level.

C Description of the data

C.1 General remarks on the retail interest rate statistics

The examined Austrian retail interest rates come from the national interest rate statistics and were compiled from 1995 until June 2003. Data source (as for the Austrian interbank rates) is the Austrian Central Bank, the Oesterreichische Nationalbank (OeNB). Rates are nominal (plus certain fees, but commissions on turnover were not included) and expressed as annual percentages.

Business coverage: Banks report the interest rate charged most frequently for new business (renewals are not considered).

Institutional coverage: Sample of 43 Monetary Financial Institutions (had decreased to 37 banks in 2003 because of mergers). "This sample consisted of the major joint stock banks, the state mortgage banks as well as the largest institutions of the savings bank, Raiffeisen credit cooperative and Volksbank credit cooperative sectors." (Klein et al. 2003)

Aggregation method: Arithmetic averages excluding 5% of the rates at both ends of the range.

Note: Interest rates on loans from home savings banks (building and loan associations), which play an important role for housing finance in Austria, were not recorded.

C.2 The interest rate series in detail

Monetary policy rate: Overnight VIBOR (Vienna Interbank Offered Rate). The Eonia (Euro Overnight Index Average, source: European Central Bank, ECB) is appended from 1999 on.

Money market (interbank) rates: 1-month, 3-month, 6-month and 12-month VIBOR, Euribor (Euro Interbank Offered Rate, source: ECB) appended.

Loans to enterprises (commercial loans): These floating-rate loans are usually short-term (up to 1 year).

Housing loans to households: All loans used for purchasing housing space which are not mortgage loans. Rates are floating/variable and housing loans are typically long-term loans.

Consumer loans (to households): Secured, but not necessarily by mortgage. Usually long-term, but no breakdown by maturity is available.

Discount loans: Short-term credit with banks purchasing trade bills.

Mortgage loans to households and enterprises (hypothecary credit). Floating-rate loans secured by a mortgage recorded in the land register, usually long-term.

Credit to public-sector authorities (municipal loans, usually long-term).

Overdrafts on cash accounts.

Current account (sight) deposits.

Overnight savings deposits (savings not being tied up).

Savings deposits with an agreed maturity of up to 12 months.

Savings deposits with an agreed maturity of over 12 months.

Secondary market yield of bonds of domestic banks with a fixed rate of interest. Average yield of all bonds with more than one year to maturity weighted by outstanding volumes (Source: Oesterreichische Kontrollbank, OeKB).