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Expected Monetary Policy and the Dynamics of Bank Lending Rates

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Editorial

In this paper the authors explore empirically to what extent expected monetary policy matters for the dynamics of bank lending rates in the U.S., the U.K. and Germany. The authors find that banks have increasingly behaved in a forward-looking fashion by taking expected changes in monetary policy rates into account when setting lending rates. They document that along with the shifts in monetary policy regimes towards inflation targeting, expected monetary policy has become more important as a determinant of bank lending rates. Overall, their results provide support for the hypothesis that monetary policy has become more effective by successfully influencing private sector expectations.

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Expected Monetary Policy and the Dynamics of Bank Lending Rates

Claudia Kwapil[*] Johann Scharler[†]

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Abstract

In this paper we explore empirically to what extent expected monetary policy matters for the dynamics of bank lending rates in the U.S., the U.K. and Germany. We find that banks have increasingly behaved in a forward-looking fashion by taking expected changes in monetary policy rates into account when setting lending rates. We document that along with the shifts in monetary policy regimes towards inflation targeting, expected monetary policy has become more important as a determinant of bank lending rates. Overall, our results provide support for the hypothesis that monetary policy has become more effective by successfully influencing private sector expectations.

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1 Introduction

We analyze empirically to what extent banks take expected monetary policy into account when setting lending rates and how the importance of expectations has changed over time. Our analysis is motivated by the idea that the effectiveness of monetary policy is closely related to the extent to which expectations of the private sector can be influenced. Goodfriend (1991) argues that monetary policy manages to influence long-term interest rates not just by adjusting the target for current short-term rates, but also - and perhaps even more importantly - by influencing the expectations of the path of future short-term rates, as this is a major determinant of long-term interest rates. Put differently, the more credible and predictable monetary policy is, the more effective it should be. Along these lines, Woodford (2003) argues that policy inertia strongly increases predictability and thereby fosters the effectiveness of monetary policy.

It appears conceivable that this argument applies not only to long-term market rates but also to the setting of bank retail rates. If retail interest rates respond not only to the current stance of monetary policy but also to expected monetary policy, then the pass-through from monetary policy rates to retail rates might be faster, as banks react to some extent even before monetary policy is fully adjusted. In addition, the overall extent to which changes in monetary policy are passed through to retail rates may be larger. Consequently, monetary policy is more effective in influencing aggregate demand as compared to the case where banks do not incorporate forecasts of future monetary policy actions into their pricing decisions.

We find that expected future changes in monetary policy rates influence bank lending rates in the U.S. and the U.K. and that the impact of expected policy changes has increased over time. For Germany, our results are not as clear-cut. Nevertheless, despite a relatively minor effect that expected policy rate changes exert on lending rates, we also find that the overall pass-through from monetary policy rates to lending rates has increased.

Overall, we conclude that banks have transmitted changes in policy rates to lending rates to a greater extent, which is at least partly due to an increased influence of expected monetary policy. Since one would expect that banks are more likely to adjust retail rates if they believe that a change in monetary policy rates, which determine the cost of holding reserves, will not be reversed for a period of time unless warranted by a change in conditions, this result is consistent with the interpretation that monetary policy is to some extent perceived as predictable and also credible.

A large literature argues that monetary policy has become well managed over time by switching to an interest rate rule which puts a sufficiently high weight on inflation (see e.g. Judd and Rudebusch, 1998; Clarida, Galí, and Gertler, 1998, 1999, 2000; Leduc, Sill, and Stark, 2007). Similarly, Assenmacher-Wesche (2006) also argues that monetary policy has been characterized by regime switches, although her results indicate that the switches occurred somewhat later. Our results suggest a complementary explanation for why monetary policy has become more stabilizing, namely, that in addition to the change in the monetary policy regime, monetary policy has become more effective over time due to a faster and stronger transmission to interest rates directly relevant for the determination of aggregate demand. In line with our interpretation, we find that the break points after which expected monetary policy is passed on to a greater extent to lending rates, correspond closely to those identified by Clarida, Galí, and Gertler (1998) and Assenmacher-Wesche (2006).

The remainder of the paper is organized as follows: Section 2 outlines the empirical model used for assessing the role of expectations in the price setting of banks. Section 3 presents our estimation results and in Section 4 we test for break points in our estimating equation. Finally, Section 5 summarizes and concludes the paper.

2 The Empirical Model

Our analysis is based on the following equation for the dynamics of the lending rate:

$$\Delta LR_t = \alpha + \beta E[(MR_{t+k} - MR_t)|\Omega_t] + \sum_{i=0}^{m} \delta_i \Delta MR_{t-i} + \sum_{j=1}^{n} \gamma_j \Delta LR_{t-j}, \tag{1}$$

where LR_t and MR_t denote the retail lending rate and the monetary policy rate, respectively. E is the expectation operator, Ω_t is the information set at time t and Δ is the difference operator. Hence, we postulate that a change in the retail lending rate at time t is determined by the expected change of the monetary policy rate between t and $t+k$, the current change in the policy rate, m lagged changes of the policy rate and n of its own lags.

Equation (1) is closely related to the equations estimated in the empirical literature on the interest rate pass-through (see e.g. De Bondt, Mojon, and Valla, 2005; De Bondt, 2005; Sander and Kleimeier, 2004; Mojon, 2000; Borio and Fritz, 1995; Cottarelli and Kourelis, 1994), which studies the extent to which retail interest rates respond to changes in market interest rates. A typical result in this literature is that retail rates are sticky with respect to money market rates. Put differently, changes in money market rates lead to a less than one-to-one change in retail interest rates. Theoretically, it is not entirely clear why retail interest rates are sticky to a certain extent. Hofmann and Mizen (2004) and Hannan and Berger (1991) argue that sticky retail interest rates may be the result of adjustment costs. Berger and Udell (1992) find evidence for the hypothesis that banks implicitly provide insurance against large movements in rates by adjusting retail rates only to a limited extent.

Additionally to what is typically included in a pass-through equation, an expectations term enters equation (1), where β measures the influence of expected changes in monetary policy rates on the change of the lending rate at time t. For $\beta = 0$, equation (1) reduces to the pass-through equation typically estimated in the literature. Thus, our formulation nests the standard pass-through equation.

The formation of expectations has been analyzed in the literature on the interest rate pass-through only to a limited extent. De Bondt, Mojon, and Valla (2005) argue that expectations might be an important element of the pass-through process and include long-term interest rates in their specification to control for expectations.[1]

Hülsewig, Mayer, and Wollmershäuser (2006) also assume that banks are forward-looking and follow a Calvo (1983) pricing scheme. They derive an expression where the dynamics of the lending rate are partially determined by expected interest rates. Hence, their model provides additional support for the inclusion of the forward-looking part in equation (1).

If adjusting retail rates is indeed costly, as in Hofmann and Mizen (2004) and Hannan

[1]A related but distinct issue is analyzed by Sander and Kleimeier (2006), who distinguish between expected and unexpected changes in monetary policy and study the implications for the pass-through to retail rates. In contrast to their paper, our analysis focuses on the degree to which banks incorporate future changes in monetary policy into their pricing decisions.

and Berger (1991), then expectations become even more important, since anticipating future changes in monetary policy and acting accordingly may help to reduce these costs. Thus, banks have a clear incentive to take expected monetary policy into account when setting retail rates. Similarly, if the primary reason for the sluggish behavior of retail interest rates is liquidity insurance, as in Berger and Udell (1992), the response of the banking sector to changes in the monetary policy stance will depend on the perceived persistence of changes in monetary policy rates. Consequently, banks have again an incentive to take expected monetary policy into account.

Under rational expectations, equation (1) implies a set of orthogonality conditions and, therefore, the model can be estimated by the generalized method of moments (GMM):

$$E\left\{(\Delta LR_t - \alpha - \beta(MR_{t+k} - MR_t) - \sum_{i=0}^{m} \delta_i \Delta MR_{t-i} - \sum_{j=1}^{n} \gamma_j \Delta LR_{t-j}) z_t\right\} = 0, \qquad (2)$$

where z_t denotes a vector of instruments known at time t, that is $z_t \in \Omega_t$.

To find suitable instruments, we assume that monetary policy rates are set according to a potentially forward-looking interest rates rule. Moreover, we assume that Ω_t is identical to the information set available to the central bank at time t. That is, when setting retail lending rates, banks know the interest rate rule according to which monetary policy rates are set, and they have the same information the central bank uses to generate forecasts concerning the future state of the economy. This assumption allows us to draw on the large literature on the estimation of monetary policy reaction functions. Following Clarida, Galí, and Gertler (1998, 2000) we use the inflation rate and the output gap to instrument the expected policy rate. In addition, we include the long-term bond yield in the instrument vector, since this variable may also contain information on the future path of monetary policy rates (De Bondt, Mojon, and Valla, 2005).

3 Data and Results

3.1 Data

To estimate equation (2) we set the forecast horizon, k, to be 1, 3 and 6 months, in order to cover a broad range of relevant horizons. The next choice confronting any empirical research on bank lending rates is the selection of a specific rate variable, LR_t, where we use the interest rate for short-term business loans from the BIS database. Despite differences in the structure of retail banking markets and in statistical systems in the U.S., the U.K. and Germany, the BIS database ensures that the series are reasonably similar across countries.

For the U.S., this series is the prime rate charged by banks on short-term business loans at the end of the month, which is posted by the Federal Reserve Bank. For the U.K. it is the base rate ("blue chip") of London clearing banks published by the Central Statistical Office plus 100 basis points. Finally, for Germany it is the monthly average interest rate charged on current account credits between 100,000 and 500,000 Euro from basically all German banks. This series is published by the Bundesbank.[2]

As proxy for the monetary policy rate, MR_t, we use the interest rate on the spot money market in the respective currency from the BIS database, which is the overnight rate.[3] For our purpose of dealing with expectations on future monetary policy separately, the overnight money market rate is preferable to the three-month money market rate, as the three-month money market rate not only mirrors the current monetary policy stance but also market expectations on future interest rates.

As instruments, we use for each country the consumer price index to measure inflation (π_t), the deviation of the logarithm of industrial production from its quadratic trend to measure the output gap (y_t), and the yield on 10-year government bonds (BR_t). All series have a monthly frequency and are from the BIS database as well. The number of lags of ΔLR_t and ΔMR_t are set to $m = n = 6$ for the U.S. and the U.K. and to $m = 4$ and $n = 6$ for Germany, which is sufficient to whiten the residuals.

3.2 Expected Monetary Policy and Bank Lending Rates

In this section we present estimation results based on equation (2) for the U.S., the U.K. and Germany. We estimate the equation for different sub samples, where the split dates are chosen according to Clarida, Galí, and Gertler (1998), to analyze if and how the interest rate pass-through has changed along with the switch in monetary policy documented in the literature. Clarida, Galí, and Gertler (1998) argue that since 1979,

[2]Note that Borio and Fritz (1995) use the same time series for investigating the pass-through of monetary policy to lending rates in these three countries.

[3]For more details on the interest rate series, see Table A1 in Appendix A.

central banks have typically pursued some form of inflation targeting and have followed a stable interest rate rule, which allows central banks to build credibility. While in the 1960s and 1970s the reactions of the central banks to changes in inflation rates were small, controlling inflation became a major focus thereafter. The break dates differ slightly from country to country, because we choose the actual institutional changes as the beginnings of the new inflation fighting regimes. For the U.S., we take October 1979, when the new chairman of the Board of Governors of the Federal Reserve System was appointed. For the U.K. our break point is June 1979, when fighting inflation became a policy objective for the Bank of England. Finally, for Germany we pick March 1979, the time the Bundesbank entered the European Exchange Rate Mechanism.

When estimating equation (2), we are especially interested in the coefficient on expected interest rate changes, β, the immediate pass-through, δ_0, and the long-run pass-through, λ, which we calculate as

$$\lambda = \frac{k\beta + \sum_{i=0}^{m} \delta_i}{1 - \sum_{j=1}^{n} \gamma_j}. \tag{3}$$

Note that this definition of the long-term pass-through takes into account past as well as future changes in policy rates, where k gives the length of the forecast horizon. Hence, if banks pass on expected changes in policy rates even before the actual monetary policy decision has been taken, this change will be incorporated in our measure of the long-term pass-through.

Tables 1, 2 and 3 display the results for the U.S., the U.K. and Germany, respectively. The role of expected monetary policy is captured by β in our analysis. For the U.S. we see from the second column of Table 1 that expected monetary policy had a significant impact on the dynamics of the lending rate only for forecast horizons of one and three months in the earlier sub sample. After the switch in monetary policy in the late 1970s, our point estimates for β are generally larger and significant regardless of the forecast horizon. Hence, after the switch in monetary policy, banks in the U.S. have responded to expected monetary policy changes to a greater extent, indicating that monetary policy has had a stronger leverage over lending rates by stirring the expectations of the banking sector. Thus, our results are consistent with the hypothesis that monetary policy has indeed become more predictable and ultimately more effective in influencing the interest rate-sensitive part of aggregate demand.

Along with the increase in β we also observe an increase in the immediate pass-through, δ_0, which went up from approximately 20 basis points before 1979 to around 55 to 72 basis points afterwards, implying that more than 50 percent of a change in the policy rate are passed on to borrowers within the same month. The stronger impact of expected changes in monetary policy rates and the higher immediate pass-through result in a higher long-run pass-through. While the long-run pass-through is clearly below one in the first sub sample, for the latter sub sample the null hypothesis of a complete long-run pass-through, that is $\lambda = 1$, is not rejected at standard levels of significance.

Overall, we find that monetary policy in the U.S. has been transmitted to lending rates more quickly and completely since the beginning of the 1980s. These results are in line with those reported in Moazzami (1999), whose specification does not explicitly take expectations of future policy rates into account. Hence, our results reveal that this finding is at least to some extent due to the forward-looking behavior of banks.

From the results presented in Table 2 for the U.K., we see that broadly similar conclusions emerge. The point estimates of the coefficients on the expected change in the monetary policy rate, β, have increased in magnitude for all forecast horizons and are highly significant in the second sub sample. Likewise, the immediate pass-through, δ_0, has increased over time with point estimates of magnitudes similar to those in the U.S. Consequently, the long-run pass-through has also increased and turns out to be complete in the second sub sample.

Turning to the estimation results for Germany shown in Table 3, we see that the U.S. and the U.K. bear a higher resemblance to each other than to Germany. Expected monetary policy had a quantitatively small though significant impact on lending rates prior to 1979. After the sample split, the null hypothesis that $\beta = 0$ is rejected only for $k = 3$. Hence, we do not find much evidence in favor of the hypothesis that monetary policy in Germany has had a larger impact on bank lending rates via its impact on the formation of expectations since the 1980s. These results are consistent with Assenmacher-Wesche (2006), who finds only limited support for a permanent regime shift in German monetary policy in the late 1970s.

However, an increase in the immediate pass-through, δ_0, is also evident in Germany. While before 1979 the pass-through within one month was negligible, it increased to approximately 15 basis points after 1979. The data for Germany do not give a clear picture on the size of the long-run pass-through, λ, before 1979. Nevertheless, the estimation results for the sample starting in 1979 are in favor of an increase in the long-run pass-through, which in the 1980s and 1990s amounts to approximately 60 basis points. Hence, the long-run pass-through is clearly incomplete in Germany in the latter sub sample, suggesting that German banks pass on changes in monetary policy rates only to a limited extent even in the long run.

Overall, changes in policy rates are passed on to firms to a lesser extent in Germany than in the U.S. and the U.K. Hannan and Berger (1991) argue that the stickiness of retail interest rates is likely to be the result of limited competition in the banking sector. Similar conclusions are drawn in Kok Sorensen and Werner (2006). Hence, our results are compatible with this interpretation, as the German banking sector is generally thought to be more regulated and less competitive (Hofmann, 2006). Moreover, long-run relationships between banks and firms that may give rise to implicit interest rate insurance appear to be particularly close in a bank-based system like the German one (Semenov, 2006).

4 Testing for Break Points

So far, we have demonstrated that economically significant changes have occurred in the interest rate pass-through processes in the U.S. and the U.K. and to a lesser extent also in Germany. To check whether these changes are also statistically significant, we apply the structural stability test for known and unknown break points suggested by Hall and Sen (1999).

4.1 Testing for Known Break Points

Based on the distinction between identifying and over-identifying restrictions, the test by Hall and Sen (1999) allows us to distinguish between the case where the instability is confined to the parameters of the model and the case where it is related to other aspects.

Table 4 shows the results. The W-test is the Wald-test statistic for testing parameter constancy. The O-test is the statistic for testing the stability of the over-identifying restrictions. Under the null hypothesis, the over-identifying restrictions are valid before and after the break point (see Hall and Sen, 1999). For all three countries the W-test and O-test unanimously suggest that the source of the instability emerges from the parameters, whereas the over-identifying restrictions are valid in both sub samples.

Furthermore, they indicate that a structural break has indeed occurred in 1979 as we assume in Section 3. However, for the six-month horizon in the U.S. neither the W-test nor the O-test rejects the null of stability. Here, the increases in the pass-through coefficients do not appear to be significant despite the fact that they are quantitatively similar to those shown in Table 1.

On balance, the break point test supports the hypothesis that banks have responded more strongly to expected changes in monetary policy since the beginning of the 1980s, which is compatible with the interpretation that central banks have indeed become more credible and predictable over time.

4.2 Testing For Unknown Break Points

Although the institutional changes in the U.S., the U.K. and Germany at the end of the 1970s and their relevance for monetary policy are well documented in the literature (see e.g. Clarida, Galí, and Gertler, 1998), considerable uncertainty remains about the precise timing and effects of these changes in the institutional environment. Therefore, we now test for unknown break points in equation (2) as an additional robustness check. That is, instead of imposing a break point, we now let the data speak and estimate the break point.

The test proposed by Hall and Sen (1999) offers an extension for the case when the break point is unknown. Based on the W-test and the O-test described in the previous section, the composite null hypothesis of stability at each point in time within a specified interval is tested. We limit our investigation to the central 70 percent of the data points of our sample, which is standard in the literature. That is, we search for break points in the period 1974 to 2000 in the U.S., 1978 to 2001 in the U.K., and 1976 to 1994 in Germany. The W-test and the O-test statistics are calculated for each point in time given the above time interval. Hall and Sen (1999) use this sequence to construct single test statistics, where they focus on three different variants, $supW_t$ (the supremum), avW_t (the average form) and $expW_t$ (the exponential form). The $supO_t$, avO_t, and $expO_t$ statistics are constructed analogously.

The results of the stability tests are given in Table 5 for the U.S. and the U.K. and for Germany in Table 6. The null hypothesis of stability based on the over-identifying restrictions is never rejected. Thus, these tests again suggest unanimously for all three countries and over all three horizons that the instability occurs in the parameters only and not in other aspects of the model. Hence, the model as well as the choice of the instruments exhibit a stable relationship regardless of the break point chosen, which confirms our previously reported results.

For the U.S., the test indicates that the most important break in the relationship incorporating expectations over one month is December 1980, while it is March 1981 when expectations over three months are taken into consideration. Hence, for the one-month and three-months horizons, our previous results based on the break dates reported in Clarida, Galí, and Gertler (1998) are confirmed. For the six-months horizon a considerably later break date is suggested, namely February 1994.

Interestingly, these results are in line with Assenmacher-Wesche (2006), who finds two shifts in U.S. monetary policy regimes, which occurred in 1980 and approximately at the end of 1993. The first switch coincides with the change in monetary policy initiated by the new chairman of the Federal Reserve and is in line with our results for the one-month and three-months horizons. Assenmacher-Wesche interprets the second shift in 1993 as the return to a low-inflation regime after the cuts in interest rates around the 1990/1991 recession. In any case, the degree to which banks have adjusted lending rates to expected changes in monetary policy has increased over time in the U.S.

The break dates for the U.K. are June 1984 (for the one-month horizon) and March and October 1978 (for the three- and six-months horizons, respectively). The break points indicated for the three- and six-months horizons are again very close to those chosen on the basis of Clarida, Galí, and Gertler (1998). For the one-month horizon, however, the test indicates a later change in the relationship. These results for the U.K. are again in line with the switch dates in Assenmacher-Wesche (2006). She finds that the U.K. was in a high-inflation regime until 1984 and shifted to a more restrictive policy regime thereafter. Moreover, she argues that the shift to the low-inflation regime coincides roughly with the disinflation starting at the beginning of the 1980s. Overall, we may conclude that the regime shift in the U.K. has been a more gradual process, which might have been initiated at the end of the 1970s, but was completed only some years later in 1984.

Finally, we test for an unknown break point with German data. Table 6 gives the results. Again, we find that the instability is caused by the parameters only and not by other aspects of the model. The break date for the one-month horizon is April 1979, which nearly coincides with our earlier assumption. However, for the three- and six-months horizons, the test suggests a later break point, namely September 1992, which is likely to be related to the breakdown of the European Monetary System (EMS) and also to the German reunification.

To see whether the choice of the break date has an effect on the coefficient estimates in the sub samples, we re-estimate equation (2) with break dates chosen according to those in Tables 5 and 6. Here, we estimate the equation only for the U.K. and Germany, since for the U.S. the information given in Table 5 together with the interpretation in Assenmacher-Wesche (2006) does not provide evidence of another fundamental change in monetary policy except the one at the end of the 1970s.

For the U.K., we split the sample in June 1984. From Table 7 we see that our conclusions remain unchanged. The immediate pass-through and the long-run pass-through are a bit higher in the first sub sample compared to the results presented in Table 2, as the sub sample extends into a more recent period. However, the finding that the role of expectations in the price-setting process of banks became more important and the result that the immediate as well as the long-run pass-through increased over time are unchanged.

For Germany we re-estimate equation (2) for two sub samples ranging from June 1972 to March 1979 and from April 1979 to September 1992, as suggested by the test for unknown break points. The results are given in Table 8. Since September 1992 lies near the end of our sample, it does not seem meaningful to estimate a third sub sample from 1992 to 1998, especially as the period covers the break down of EMS and the related turbulence. Again, the results are very similar to those in Table 3 and the main message remains unaltered. Expected monetary policy does not seem to matter much in Germany, neither before 1979 nor afterwards. However, in the second sub sample, which covers the EMS period, the immediate pass-through as well as the long-run pass-through increased compared to the time before the EMS.

Overall, we conclude that the analysis of unknown break points conducted in this section confirms our previous results. Although some uncertainty remains concerning the point at which the most fundamental change in the relationship between expected monetary policy and bank lending rates has occurred, our main conclusions remain unaltered. Banks have incorporated forecasts of future monetary policy to a greater extent over time in the U.S. and the U.K. and the overall pass-through has increased in all three countries in our sample.

For the U.S. the test for an unknown break point confirms that the most relevant change in the transmission of monetary policy to bank lending rates has occurred in the late 1970s, which is in line with the switch to a stabilizing interest rule documented in Clarida, Galí, and Gertler (1998). For the U.K. the test indicates an additional, somewhat later break point, namely 1984. For Germany, we find that the breakdown of the EMS in 1992 may have had rather substantial consequences for the predictability of monetary policy.

5 Concluding Remarks

In this paper we analyze the extent to which expected monetary policy matters for the dynamics of retail interest rates. We find that expected changes in monetary policy rates influence bank lending rates in the U.S. and the U.K. and to a limited extent also in Germany. Moreover, we find that banks have become more forward-looking over time in the U.S. and in the U.K., which is consistent with the hypothesis that monetary policy has become more predictable along with the implementation of a stabilizing interest rate rule. Consequently, monetary policy is transmitted faster and to a greater extent to lending rates. Put differently, monetary policy has had a growing impact on retail interest rates via its ability to influence expectations of the private sector as described in Goodfriend (1991), Woodford (2003), and Galí and Gertler (2007).

We provide a complementary explanation for the higher macroeconomic stability observed over the past decades. Clarida, Galí, and Gertler (1998, 1999, 2000) argue that monetary policy has been relatively successful since the late 1970s mostly because central banks have responded sufficiently to increases in inflation to rule out self-fulfilling expectations. We add that higher macroeconomic stability has also been the result of the growing impact of monetary policy on interest rates relevant for aggregate demand.

Some caveats have to be kept in mind. It appears conceivable that the increase in the overall pass-through and also in the forward-looking behavior of banks are not primarily related to the predictability of monetary policy. One might also argue that the fact that banking markets have become less regulated and more competitive over time results in an increase in the pass-trough. However, our break points correspond closely to regime shifts in monetary policy, suggesting that the increase in the forward-looking behavior is indeed related to monetary policy.

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A Appendix

Table A1: Data Description of Money Market and Retail Interest Rates

	Source	Codes	Time Period
United States			
Prime/base rate on short-term loans	BIS	HLBAUS01	1968:08 - 2006:12
Money market rate, overnight	BIS	JBBAUS02	1968:08 - 2006:12
United Kingdom			
Prime/base rate on short-term loans	BIS	HLBAGB11	1973:08 - 2006:12
Money market rate, overnight	BIS	JBBAGB02	1973:08 - 2006:12
Germany			
Prime/base rate on short-term loans	BIS	HLBADE06	1972:06 - 1998:12
Money market rate, overnight	BIS	JBBADE91	1972:06 - 1998:12

Notes: BIS stands for the Data Base of the Bank for International Settlements.

Table 1: Pass-Through to Lending Rates in the U.S.

	1968:08 - 1979:09						1979:10 - 2006:12						
k	β		δ_0		λ		β		δ_0		λ		
1 month	0.20	***	0.22	***	0.69	***	0.35	***	0.55	***	1.13	***	$R^2 = 0.70$
	(0.05)		(0.02)		(0.06)		(0.05)		(0.04)		(0.06)		$T = 460$
3 months	0.04	**	0.15	***	0.45	***	0.04	**	0.72	***	1.01	***	$R^2 = 0.72$
	(0.02)		(0.03)		(0.10)		(0.02)		(0.02)		(0.05)		$T = 458$
6 months	0.02		0.21	***	0.78	***	0.06	***	0.66	***	1.06	***	$R^2 = 0.70$
	(0.01)		(0.03)		(0.06)		(0.01)		(0.03)		(0.05)		$T = 455$

Notes: The standard errors are shown in parenthesis. In our baseline model for the U.S. the instrument vector z_t includes a constant, ΔMR_t, ΔMR_{t-1}, ΔMR_{t-2}, ΔMR_{t-3}, ΔMR_{t-4}, ΔMR_{t-5}, ΔMR_{t-6}, ΔLR_{t-1}, ΔLR_{t-2}, ΔLR_{t-3}, ΔLR_{t-4}, ΔLR_{t-5}, ΔLR_{t-6}, Δy_{t-1}, Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, Δy_{t-5}, Δy_{t-6}, Δy_{t-9}, Δy_{t-12}, $\Delta \pi_{t-1}$, $\Delta \pi_{t-2}$, $\Delta \pi_{t-3}$, $\Delta \pi_{t-4}$, $\Delta \pi_{t-5}$, $\Delta \pi_{t-6}$, $\Delta \pi_{t-9}$, $\Delta \pi_{t-12}$, ΔBR_{t-1}, ΔBR_{t-2}, ΔBR_{t-3}, ΔBR_{t-4}, ΔBR_{t-5}, ΔBR_{t-6}, ΔBR_{t-9}, ΔBR_{-12}. The J-tests (see Hansen, 1982) do not reject the null hypothesis of validity of the instruments. The standard errors for the long-term pass-through are calculated according to the delta method (see e.g. Greene, 2000, p. 330).

Table 2: Pass-Through to Lending Rates in the U.K.

	1973:08 - 1979:05						1979:06 - 2006:12						
k	β		δ_0		λ		β		δ_0		λ		
1 month	0.09	***	0.14	***	0.34	***	0.99	***	0.50	***	1.00	***	$R^2 = 0.39$
	(0.03)		(0.02)		(0.02)		(0.10)		(0.05)		(0.03)		$T = 396$
3 months	0.08	***	0.26	***	0.78	***	0.34	***	0.55	***	1.09	***	$R^2 = 0.39$
	(0.02)		(0.03)		(0.07)		(0.05)		(0.07)		(0.05)		$T = 396$
6 months	-0.01		0.20	***	0.52	***	0.27	***	0.73	***	1.40	***	$R^2 = 0.27$
	(0.01)		(0.03)		(0.08)		(0.05)		(0.06)		(0.18)		$T = 395$

Notes: The standard errors are shown in parenthesis. In our baseline model for the U.K. the instrument vector z_t includes a constant, ΔMR_t, ΔMR_{t-1}, ΔMR_{t-2}, ΔMR_{t-3}, ΔMR_{t-4}, ΔMR_{t-5}, ΔMR_{t-6}, ΔLR_{t-1}, ΔLR_{t-2}, ΔLR_{t-3}, ΔLR_{t-4}, ΔLR_{t-5}, ΔLR_{t-6}, Δy_{t-1}, Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, Δy_{t-5}, Δy_{t-6}, Δy_{t-9}, Δy_{t-12}, $\Delta\pi_{t-1}$, $\Delta\pi_{t-2}$, $\Delta\pi_{t-3}$, $\Delta\pi_{t-4}$, $\Delta\pi_{t-5}$, $\Delta\pi_{t-6}$, $\Delta\pi_{t-9}$, $\Delta\pi_{t-12}$, ΔBR_{t-1}, ΔBR_{t-2}, ΔBR_{t-3}, ΔBR_{t-4}, ΔBR_{t-5}, ΔBR_{t-6}, ΔBR_{t-9}, ΔBR_{-12}. The J-tests (see Hansen, 1982) do not reject the null hypothesis of validity of the instruments. The standard errors for the long-term pass-through are calculated according to the delta method (see e.g. Greene, 2000, p. 330).

Table 3: Pass-Through to Lending Rates in Germany

	1972:06 - 1979:02						1979:03 - 1998:12						
k	β		δ_0		λ		β		δ_0		λ		
1 month	0.01	**	0.01		0.25		-0.07		0.18	***	0.56	***	$R^2 = 0.46$
	(0.00)		(0.01)		(0.24)		(0.06)		(0.02)		(0.05)		$T = 319$
3 months	0.01	**	0.00		-0.04		0.03	*	0.13	***	0.58	***	$R^2 = 0.47$
	(0.00)		(0.00)		(0.08)		(0.02)		(0.02)		(0.05)		$T = 319$
6 months	0.02	***	0.02	***	0.62	***	0.01		0.14	***	0.57	***	$R^2 = 0.46$
	(0.00)		(0.00)		(0.05)		(0.02)		(0.02)		(0.07)		$T = 319$

Notes: The standard errors are shown in parenthesis. In our baseline model for Germany the instrument vector z_t includes a constant, ΔMR_t, ΔMR_{t-1}, ΔMR_{t-2}, ΔMR_{t-3}, ΔMR_{t-4}, ΔLR_{t-1}, ΔLR_{t-2}, ΔLR_{t-3}, ΔLR_{t-4}, ΔLR_{t-5}, ΔLR_{t-6}, Δy_{t-1}, Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, Δy_{t-5}, Δy_{t-6}, Δy_{t-9}, Δy_{t-12}, $\Delta\pi_{t-1}$, $\Delta\pi_{t-2}$, $\Delta\pi_{t-3}$, $\Delta\pi_{t-4}$, $\Delta\pi_{t-5}$, $\Delta\pi_{t-6}$, $\Delta\pi_{t-9}$, $\Delta\pi_{t-12}$, ΔBR_{t-1}, ΔBR_{t-2}, ΔBR_{t-3}, ΔBR_{t-4}, ΔBR_{t-5}, ΔBR_{t-6}, ΔBR_{t-9}, ΔBR_{-12}. The J-tests (see Hansen, 1982) do not reject the null hypothesis of validity of the instruments. The standard errors for the long-term pass-through are calculated according to the delta method (see e.g. Greene, 2000, p. 330).

Table 4: Tests for Structural Breaks when the Break Point Is Known

	1 month		3 months		6 months	
U.S. (Oct. 1979)						
W-test	19.16	**	17.15	**	8.12	
O-test	21.09		15.58		13.69	
U.K. (June 1979)						
W-test	33.79	***	47.66	***	37.76	***
O-test	16.18		14.53		12.24	
Germany (March 1979)						
W-test	102.84	***	99.95	***	115.21	***
O-test	19.53		16.87		18.49	

Notes: The break points for each country are given in parenthesis. Andrews and Fair (1988) show that the W-test statistic converges to a χ^2_p distribution under the null of parameter constancy. Following Hall and Sen (1999), the O-test statistic converges to a χ^2_{2q-2p} distribution. The number of coefficients we are estimating, p, is 9 (α, β, γ_1, γ_2, γ_3, δ_0, δ_1, δ_2, δ_3). The number of instruments, q, we are using for GMM is 17 (constant, ΔMR_t, ΔMR_{t-1}, ΔMR_{t-2}, ΔMR_{t-3}, ΔLR_{t-1}, ΔLR_{t-2}, ΔLR_{t-3}, Δy_{t-1}, Δy_{t-2}, Δy_{t-3}, $\Delta\pi_{t-1}$, $\Delta\pi_{t-2}$, $\Delta\pi_{t-3}$, ΔBR_{t-1}, ΔBR_{t-2}, ΔBR_{t-3}).

Table 5: Tests for Structural Breaks when the Break Point Is Unknown

	1 month		3 months		6 months		Critical Values at 1%
U.S.							
$supW_t$	83.57	***	90.64	***	69.45	***	30.42
avW_t	37.60	***	36.86	***	40.13	***	17.30
$expW_t$	36.03	***	39.57	***	30.20	***	11.67
$supO_t$	22.08		23.55		22.01		40.09
avO_t	17.18		16.95		17.17		29.52
$expO_t$	9.10		8.93		9.04		16.80
break point	Dec. 1980		March 1981		Feb. 1994		
U.K.							
$supW_t$	100.84	***	92.35	***	74.95	***	30.42
avW_t	34.16	***	26.19	***	23.76	***	17.30
$expW_t$	44.87	***	40.58	***	32.39	***	11.67
$supO_t$	22.66		17.07		18.81		40.09
avO_t	17.41		14.59		14.97		29.52
$expO_t$	9.25		7.44		8.02		16.80
break point	June 1984		March 1978		Oct. 1978		

Notes: The tests for unknown break points start after the first and end before the last 15% of observations ($\Pi = 0.15, 0.85$). The number of coefficients we are estimating, p, is 9 (α, β, γ_1, γ_2, γ_3, δ_0, δ_1, δ_2, δ_3). The number of instruments, q, we are using for GMM is 17 (constant, ΔMR_t, ΔMR_{t-1}, ΔMR_{t-2}, ΔMR_{t-3}, ΔLR_{t-1}, ΔLR_{t-2}, ΔLR_{t-3}, Δy_{t-1}, Δy_{t-2}, Δy_{t-3}, $\Delta\pi_{t-1}$, $\Delta\pi_{t-2}$, $\Delta\pi_{t-3}$, ΔBR_{t-1}, ΔBR_{t-2}, ΔBR_{t-3}). For $supW_t$ the critical values are taken from Table 1 in Andrews (1993) and Andrews (2003) with $p = 9$. For avW_t and $expW_t$ the critical values are taken from Table 2 and 1, respectively, in Andrews and Ploberger (1994) with $p = 9$. For $supO_t$, avO_t and $expO_t$ the critical values are taken from Table 1 in Hall and Sen (1999) with $q - p = 8$. The break dates reported in this Table are chosen according to $supW_t$.

Table 6: Tests for Structural Breaks when the Break Point Is Unknown

	1 month		3 months		6 months		Critical Values at 1%
Germany							
$supW_t$	121.66	***	176.98	***	210.55	***	30.42
avW_t	59.17	***	53.52	***	82.56	***	17.30
$expW_t$	55.73	***	83.07	***	99.86	***	11.67
$supO_t$	22.40		20.56		21.80		40.09
avO_t	15.53		14.80		17.70		29.52
$expO_t$	8.83		7.96		9.15		16.80
break point	April 1979		Sep. 1992		Sep. 1992		

Notes: The tests for unknown break points start after the first and end before the last 15% of observations ($\Pi = 0.15, 0.85$). The number of coefficients we are estimating, p, is 9 (α, β, γ_1, γ_2, γ_3, δ_0, δ_1, δ_2, δ_3). The number of instruments, q, we are using for GMM is 17 (constant, ΔMR_t, ΔMR_{t-1}, ΔMR_{t-2}, ΔMR_{t-3}, ΔLR_{t-1}, ΔLR_{t-2}, ΔLR_{t-3}, Δy_{t-1}, Δy_{t-2}, Δy_{t-3}, $\Delta\pi_{t-1}$, $\Delta\pi_{t-2}$, $\Delta\pi_{t-3}$, ΔBR_{t-1}, ΔBR_{t-2}, ΔBR_{t-3}). For $supW_t$ the critical values are taken from Table 1 in Andrews (1993) and Andrews (2003) with $p = 9$. For avW_t and $expW_t$ the critical values are taken from Tables 2 and 1, respectively, in Andrews and Ploberger (1994) with $p = 9$. For $supO_t$, avO_t and $expO_t$ the critical values are taken from Table 1 in Hall and Sen (1999) with $q - p = 8$. The break dates reported in this Table are chosen according to $supW_t$.

Table 7: Pass-Through to Lending Rates in the U.K. (Alternative Break Point)

	1973:08 - 1984:05					1984:06 - 2006:12						
k	β	δ_0		λ		β		δ_0		λ		
1 month	0.02	0.42	***	0.60	***	0.73	***	0.65	***	0.93	***	$R^2 = 0.42$
	(0.05)	(0.05)		(0.03)		(0.08)		(0.05)		(0.03)		$T = 397$
3 months	-0.05	0.48	***	0.55	***	0.47	***	0.63	***	1.18	***	$R^2 = 0.30$
	(0.03)	(0.05)		(0.04)		(0.05)		(0.07)		(0.04)		$T = 397$
6 months	-0.04	0.47	***	0.53	***	0.21	***	0.56	***	1.19	***	$R^2 = 0.32$
	(0.03)	(0.05)		(0.06)		(0.03)		(0.05)		(0.10)		$T = 395$

Notes: The standard errors are shown in parenthesis. In our baseline model for the U.K. the instrument vector z_t includes a constant, ΔMR_t, ΔMR_{t-1}, ΔMR_{t-2}, ΔMR_{t-3}, ΔMR_{t-4}, ΔMR_{t-5}, ΔMR_{t-6}, ΔLR_{t-1}, ΔLR_{t-2}, ΔLR_{t-3}, ΔLR_{t-4}, ΔLR_{t-5}, ΔLR_{t-6}, Δy_{t-1}, Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, Δy_{t-5}, Δy_{t-6}, Δy_{t-9}, Δy_{t-12}, $\Delta\pi_{t-1}$, $\Delta\pi_{t-2}$, $\Delta\pi_{t-3}$, $\Delta\pi_{t-4}$, $\Delta\pi_{t-5}$, $\Delta\pi_{t-6}$, $\Delta\pi_{t-9}$, $\Delta\pi_{t-12}$, ΔBR_{t-1}, ΔBR_{t-2}, ΔBR_{t-3}, ΔBR_{t-4}, ΔBR_{t-5}, ΔBR_{t-6}, ΔBR_{t-9}, ΔBR_{-12}. The J-tests (see Hansen, 1982) do not reject the null hypothesis of validity of the instruments. The standard errors for the long-term pass-through are calculated according to the delta method (see e.g. Greene, 2000, p. 330).

Table 8: Pass-Through to Lending Rates in Germany (Alternative Break Point)

	1972:06 - 1979:03						1979:04 - 1992:09						
k	β		δ_0		λ		β		δ_0		λ		
1 month	0.01	***	0.01	**	0.40		-0.06		0.16	***	0.65	***	$R^2 = 0.48$
	(0.00)		(0.01)		(0.26)		(0.05)		(0.02)		(0.04)		$T = 244$
3 months	0.01	***	0.00		0.02		0.04	**	0.09	***	0.65	***	$R^2 = 0.45$
	(0.00)		(0.00)		(0.08)		(0.02)		(0.01)		(0.08)		$T = 244$
6 months	0.01	***	0.01	*	0.63	***	-0.01		0.14	***	0.61	***	$R^2 = 0.50$
	(0.00)		(0.00)		(0.13)		(0.02)		(0.02)		(0.05)		$T = 244$

Notes: The standard errors are shown in parenthesis. In our baseline model for Germany the instrument vector z_t includes a constant, ΔMR_t, ΔMR_{t-1}, ΔMR_{t-2}, ΔMR_{t-3}, ΔMR_{t-4}, ΔMR_{t-5}, ΔMR_{t-6}, ΔLR_{t-1}, ΔLR_{t-2}, ΔLR_{t-3}, ΔLR_{t-4}, ΔLR_{t-5}, ΔLR_{t-6}, Δy_{t-1}, Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, Δy_{t-5}, Δy_{t-6}, Δy_{t-9}, Δy_{t-12}, $\Delta \pi_{t-1}$, $\Delta \pi_{t-2}$, $\Delta \pi_{t-3}$, $\Delta \pi_{t-4}$, $\Delta \pi_{t-5}$, $\Delta \pi_{t-6}$, $\Delta \pi_{t-9}$, $\Delta \pi_{t-12}$, ΔBR_{t-1}, ΔBR_{t-2}, ΔBR_{t-3}, ΔBR_{t-4}, ΔBR_{t-5}, ΔBR_{t-6}, ΔBR_{t-9}, ΔBR_{-12}. This instrument set deviates slightly from the regressions presented in Table 3, as a higher R^2 could be achieved with this set. The J-tests (see Hansen, 1982) do not reject the null hypothesis of validity of the instruments. The standard errors for the long-term pass-through are calculated according to the delta method (see e.g. Greene, 2000, p. 330).

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