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Are Prime Rate Changes Asymmetric?

Michael J. Dueker

INTRODUCTION

Many observers have suggested that banks adjust their administered rates asymmetrically, raising loan rates more readily than they lower them. Empirical studies have found support for this claim: Depository institutions appear to move faster to lower deposit rates than to raise them (Neumark and Sharpe, 1992; Diebold and Sharpe, 1990; Hutchison,1995). Such asymmetries, if they exist in the banking system, are of interest to monetary policymakers because they would support Cover's (1992) argument that monetary policy changes are propagated asymmetrically, with monetary policy tightenings carrying more force than easings of the same magnitude. In addition, because asymmetry in loan interest rates tilted toward slow downward adjustment works to the disadvantage of loan customers, bank-dependent firms as a group are thought to be more vulnerable to the business cycle than firms with direct access to capital markets. Interest rates move in a highly procyclical manner, so that if loan rates are sluggish when moving down, relative to market interest rates, then firms that rely on bank loans will be paying relatively high rates during cyclical downturns, precisely at the time when sales revenues tend to be low. This article tests for the existence of an asymmetric relation between the prime rate and market interest rates and examines whether the effects of such asymmetry are of sufficient magnitude and duration to be of concern.

The focus of this article is the bank prime rate, a posted lending rate that ostensibly represents the rate banks offer to their most creditworthy borrowers. It is worth noting that numerous borrowers are able to get loans at less than the prime rate—especially on short-term loans—but, more importantly, many loan rates are tied to the prime, changing by an equal amount whenever the prime rate changes. Because a variety of loan rates—including many credit card balances—are tied contractually to the prime rate, it is an important benchmark. This article reviews how risk aversion on the part of bank managers could lead to asymmetric movement in bank lending rates, such as the prime. Risk aversion, in this context, simply means that managers prefer a smooth profit stream to a volatile one.

The article's main contribution, however, lies in the empirical methodology used to test for asymmetry in the response of the prime rate to short-term market interest rates. The prime rate changes discretely by multiples of 25 basis points, which implies that changes are relatively infrequent (less than 20 percent of all weeks) and, therefore, newsworthy. Discreteness can confound tests for symmetry, that is, tests for equal responsiveness to pressure to increase or decrease. The pressure to respond to movements in market interest rates does not appear suddenly only in weeks when the prime rate actually changes, so one must find a way to measure how the continuous pressure translates into occasional discrete changes. Empirical analysis of the historical data on the prime rate is hampered because econometric models provide methods for dealing with discrete-valued data and methods for analyzing time-series data, but rarely present methods for confronting the two together.

This article addresses the disparate features of the data by employing a dynamic ordered probit model of changes in the prime rate to test for two types of possible asymmetry in the prime. First, the model includes estimated thresholds at which money-center banks change the prevailing prime rate. The thresholds indicate the degree of misalignment between the prime rate and market interest rates needed to induce a discrete change in the prime rate. Second, I estimate separate coefficients for increases and decreases in the short-term interest rate. The difference between these sets of coefficients measures any asymmetry in the generation of underlying pressure for a prime rate change, whereas differences among the threshold coefficients reflect an asymmetry in how such pressure is translated into discrete prime rate changes.

WHY STUDY THE PRIME RATE?

The prime rate, not a deposit rate, is the focus in this article because the prevailing prime rate is much more uniform across banks than any measure of the prevailing deposit rate. Aggregate data for deposit rates represent an average of different rates across many banks, and such an average does not fully reflect the asymmetry that a customer might experience at a single bank. Moreover, if one were to follow the movement of a deposit rate at a single bank, its degree of asymmetry might not be representative of conditions in the whole banking system. In contrast, the prevailing prime rate among money-center banks is a figure widely quoted in the financial press, and it is generally the same across banks.

Reason to Expect Asymmetric Behavior in the Prime Rate

The idea that significant costs to customers of switching suppliers can alter the nature of competition among suppliers has found applications in industrial organization, macroeconomics, and international trade (Klemperer, 1995). A key implication is that switching costs make sellers imperfect competitors. An important example where switching costs are thought to pertain is the relationship between banks and their loan customers. Here I outline why switching costs, combined with risk-averse bank management, might lead to asymmetric movement in the pricing of bank loans.

Banks specialize in acquiring costly information about their business loan customers. Consequently, borrowers find it costly to switch from a lender who knows them to one who does not. Once a relationship is established, one might conclude that a bank could extract monopoly rents from its customers in the form of above-normal interest rates. Rajan (1992) argues that such opportunistic behavior may not fit into a bank's optimal long-run strategy, because rival banks could capture its customers by sharing the switching costs. Gilbert and Klemperer (2000) discuss ways in which competition among sellers leads to cooperation among buyers and sellers to mitigate the effects of switching costs.

Contracts in which sellers precommit to prices that compensate (at least partially) for start-up or switching costs is one such form of cooperation. For example, business loans and lines of credit are often contractually tied to the prime lending rate, the London Inter-Bank Offering Rate (LIBOR) or other cost-of-funds indexes. By tying loan rates to such indexes, banks effectively precommit to prices that are state dependent, where the state is the prevailing index rate.

Banks choose their own prime rates, however, and banks jockey to be among the first to adjust their prime rates, yet wish to avoid false starts and retractions that occur when other banks do not follow their lead. Imperfect competition rendered by switching costs leads a bank to consider the trade-off between enhancing its market share and monopoly pricing of its existing customer base when adjusting the prime rate. Several authors, including Chevalier and Scharfstein (1996) and Klemperer (1995), have observed that the business cycle can affect this trade-off if firms prefer smooth profit streams. In cyclical downturns, firms with market power may smooth profits by charging relatively high prices, rather than seeking to expand market share. Banks also must consider that if they were to seek greater market share in a cyclical downturn, they would face adverse selection: the prospect of lending to businesses with the highest cyclical probabilities of failure. Thus, the prospect of adverse selection serves as an additional source of profit volatility for banks across the business cycle. For these reasons, bank managers typically opt to maintain relatively high loan rates during cyclical downturns, rather than garner greater market share. If bank managers generally behave this way, the prime rate could display an asymmetric response to procyclical movements in short-term market interest rates. The next section discusses the empirical model used for this investigation.

Dynamic Ordered Probit Model

At their inception, qualitative response models—logits and probits—were designed for cross-sectional data (Goldberger, 1964). For probit models of a cross section, a maintained assumption is that the shock to each individual is an independent draw from the population distribution. Increasingly, however, qualitative response models are applied to time series, where one realistically cannot assume independence across observations. Discrete changes in the prime rate across time are one such example. Most qualitative response models, however, do not incorporate time-series features, such as treatment of serial correlation and conditional heteroscedasticity. The chief obstacle to applying time-series methods has been that the residuals are not readily recovered in discrete choice models.

A notable exception is the dynamic ordered probit of Eichengreen, Watson, and Grossman (1985). In this model, an observed variable, Y, changes each period by one of J different discrete amounts, including changes of zero. The impetus for a change in the prime rate does not appear suddenly in periods when discrete changes take place, only to disappear in periods when no change takes place. Instead, the dynamic probit uses a continuous measure of the impetus for a prime rate change wherein a discrete change takes place when the impetus reaches a critical level or threshold. The impetus equals the difference between a continuous latent level and the lagged value of the observed rate, $Y_t^* - Y_{t-1}$. The latent level is defined in terms of its own changes from period to period, plus an initial level, Y_0^*, where the changes in the latent level are assumed to depend on a vector of lagged explanatory variables, ΔX, which are changes in the three-month Treasury bill rate, plus a disturbance as in an ordinary regression model:

$$\Delta Y_t^* = \Delta X_{t-1}'\beta + \varepsilon_t. \tag{1}$$

The impetus is denoted Z_t and can be written as the sum of the new pressure for a change and the pressure carried from the past,

$$Z_t = \Delta Y_t^* + (Y_{t-1}^* - Y_{t-1}). \tag{2}$$

In this way, the dynamic probit model allows the impetus for a change in the prime rate to build across time periods. The size of the discrete change in the prime rate (including zero changes) depends on how the impetus compares with a set of threshold constants (c_0,..,c_j), which determine that the actual change, ΔY_t, is in category j if and only if

$$c_{j-1} < Z_t < c_j. \tag{3}$$

For example, the range for the category of no change in the prime rate might be the interval (-0.30, 0.25). This hypothetical range indicates that the gap between the latent and lagged actual prime rates needed to induce a change would have to be at least 30 basis points for a decrease and 25 basis points for an increase. In this way, the threshold coefficients can imply a form of asymmetry in the behavior of the prime rate.

The maximum-likelihood estimation procedure of Eichengreen, Watson, and Grossman (1985) for the dynamic probit requires numerical evaluation of an integral at each observation to obtain the marginal likelihood of Y_t^* from the joint density of Y_t^* and Y_{t-1}^*. In cases like the dynamic ordered probit, where the likelihood function is difficult to evaluate, Dueker (1999) shows that Bayesian analysis via a technique called Gibbs sampling offers a tractable method to generate a sample of draws from the marginal distribution of Y_t^* through a sequence of draws from the respective conditional distributions, $f(Y_t^* \mid Y_{t-1}^*, Y_{t+1}^*)$. The advantage is that one conditions the density Y_t^* on a *value*, instead of a *density*, of Y_{t-1}^*, making the problem much simpler. Further discussion of Gibbs sampling is in the accompanying insert.

Additional Features of the Model

I include Markov switching in two parameters to confront as many time-series properties of the data as possible, particularly the large increase in interest rate volatility that accompanied the Federal Reserve's use of nonborrowed reserves targets between 1979 and 1982. With Markov switching, Gibbs sampling becomes the only way to estimate the dynamic ordered probit model, because the maximum-likelihood procedure would be intractable. Markov switching in the variance of the disturbance term, ε_t, allows for time variation in the variance and should allow the model to capture the volatile 1979-82 regime. A binary state variable, $S1$, governs the state switching in the variance:

$$\begin{aligned} var(\varepsilon_t) &= \sigma_{S1_t}^2 = \sigma_0^2(1 - S1_t) + \sigma_1^2 S1_t \\ &\sigma_0^2 < \sigma_1^2 \end{aligned} \tag{4}$$

$S1_t$ is either 0 or 1.

I also make the intercept, β_0, subject to Markov switching governed by a second binary state variable, $S2$, because β_0, represents a rate of drift in Y^* since equation 1 is written in first differences. Hence, the model allows the drift parameter to shift and even to change sign. Small changes in the probability that an interest rate will increase next period can have a large impact on the prices of interest rate options, for example, so careful estimates of drift terms are important to some market participants. I also experimented with specifications that included ρY_{t-1}^* on the right-hand side of equation 1, to allow for mean reversion in the prime rate, but the estimated values of ρ were always extremely close to zero.

GIBBS SAMPLING AND THE DYNAMIC ORDERED PROBIT

In empirical estimation, it is common to encounter cases where one does not know how, or finds it cumbersome, to evaluate a probability density function or other population characteristic of a random variable of interest. For Bayesian estimation, the random variables of interest are the model parameters, such as slope coefficients (β from equation 1). We would like to make Bayesian inferences concerning the posterior distribution of β conditional on the observed data, Y_T, from the entire sample period. Such inferences would be very difficult to derive directly.

In this case, however, a conditional distribution for β is known, conditional on the latent desired levels, Y_T^*. With the Gibbs sampler, we can generate samples from $f(\beta \mid Y_T)$ using the Gibbs sampler, without ever evaluating $f(\beta \mid Y_T)$ directly (Casella and George, 1992).

The Gibbs sampler consists of a chain of draws from conditional distributions:

$$f\left(\beta^{(i+1)} \middle| \, Y_T^{*(i)}, S1_T^{(i)}, S2_T^{(i)}, Y_T\right),$$

$$f\left(Y_T^{*(i+1)} \middle| \, \beta^{(i+1)}, S1_T^{(i)}, S2_T^{(i)}, Y_T\right),$$

$$f\left(S1_T^{(i+1)}, S2_T^{(i+1)} \middle| \, \beta^{(i+1)}, Y_T^{*(i+1)}, Y_T\right),$$

where superscript i indicates iteration number i through the Gibbs sampler. That is, we draw values of β from a distribution we know (conditional on Y_T^*), and we then use those values of β to draw new values of Y_T^*. The key idea behind Gibbs sampling is that after a sufficient number of iterations of drawing new values of β and Y_T^* in this way, the draws represent draws from the respective marginal densities, conditional only on Y_T, even though the marginal densities often cannot be evaluated directly (Gelfand and Smith, 1990). As outlined in Dueker (1999), Gibbs sampling is an attractive approach for the dynamic ordered probit, because it is relatively easy to sample from appropriate conditional distributions. For β conditional on Y_T^*, the most intuitive choice for the conditional distribution is multivariate normal with the mean and variance of the ordinary least squares estimator from equation 1. For Y_T^*, the vector of latent variables, the implied conditional distribution is truncated normal. Details concerning the conditional means and variances of the elements of Y_T^* are in Dueker (1999).

The transition probabilities for the two state variables are

$$\text{Prob}\left(S1_t = 0 \middle| \, S1_{t-1} = 0\right) = p_1$$

$$\text{Prob}\left(S1_t = 1 \middle| \, S1_{t-1} = 1\right) = q_1$$

$$\text{Prob}\left(S2_t = 0 \middle| \, S2_{t-1} = 0\right) = p_2$$

$$\text{Prob}\left(S2_t = 1 \middle| \, S2_{t-1} = 1\right) = q_2.$$

Conditional distributions for the Markov state variables, $S1$ and $S2$, are included in the chain of conditional distributions for the Gibbs sampler. The conditional distributions for Markov state variables are found in Dueker (1999). Also included in the parameter chain is the vector of cut-off coefficients from equation 3. These coefficients have conditional distributions (conditional on drawn values of Z) where c_j distributed is uniformly on the interval bounded below by c_{j-1} and above by the smallest Z which is in category $j+1$. These threshold bounds ensure that $c_{j-1} < Z_t < c_j$ for all drawn values of Z.

Because it is impossible to identify simultaneously the magnitudes of the cut-off coefficients and the variances of the disturbance term, $\sigma^2_{S1_t}$, I fixed the variances at 0.01 and 0.10, respectively, for $S1 = 0$ and $S1 = 1$. This choice, in principle, allows for a tenfold shift in the variance during the Federal Reserve's period of targeting non-borrowed reserves from 1979 to 1982. In practice, the variance will not have to shift that much because the Treasury bill rate as an explanatory variable induces much of the increased volatility experienced by the prime rate. The results are robust, however, to changes in the variance levels that move them closer together, provided that the unconditional variance remains roughly the same.

Data and Estimation Results

Changes in the week-ending prime rate from January 4, 1975, to October 15, 1999, fall into seven categories with the relative frequencies shown in Table 1. The prime rate changed in about 17 percent of all weeks. I test the prime rate for asymmetric movements relative to the three-month Treasury bill rate, a benchmark short-term market interest rate. Figure 1 illustrates the procyclical swings in the T-bill rate and the prime rate. At this point, it is important to note that since April 1994, the prime rate has moved in tandem with the target federal funds rate and has been equal to the target funds rate plus 3 percent. In part, this close connection between the prime rate and the target federal funds rate owes to the practice adopted in February 1994 whereby the Federal Open Market Committee, the monetary policy body of the Federal Reserve, announces changes in the target federal funds rate at the conclusion of its meetings. Because of this regime change, which has resulted in close ties between the prime rate and the target federal funds rate, I present separate estimates for a data set that stops in April 1994 and another that continues until October 1999.[1]

Because adjustment of the prime rate to market developments may take a few weeks, we are interested primarily in tracking the cumulative response of the prime rate to changes in the market interest rate. One manifestation of asymmetry would take the form of a faster response of the prime rate to increases than decreases in the T-bill rate. Accordingly, coefficients representing sums of lag coefficients are presented. Thus, if the original model were written with Γ*s* denoting coefficients on individual lags of the regressor variable:

$$\Gamma_1 X_{t-1} + \Gamma_2 X_{t-2} + \ldots + \Gamma_k X_{t-k},$$

then the reported β coefficients are defined as

$$\beta_1 = \left(\Gamma_1 + \ldots + \Gamma_k\right), \beta_2 = \left(\Gamma_2 + \ldots + \Gamma_k\right), \ldots, \beta_k = \Gamma_k.$$

The lag length, k, was chosen based on when the β coefficients lost significance, which is at lag eight for changes in the T-bill rate. The T-bill changes are all lagged at least one period, so they are predetermined relative to this week's change in the dependent variable. I partition the T-bill changes into increases and decreases to test for asymmetric response of the latent prime rate, Y^*.

I also test for symmetry in the threshold coefficients. Equation 5 shows that symmetry would imply that $c_1 = -c_6$, $c_2 = -c_5$, and $c_3 = -c_4$. Since $Z_t = Y_t^* - Y_{t-1}$ measures the misalignment of the prime rate relative to the latent level, these restrictions on the threshold coefficients would imply that the necessary degrees of misalignment

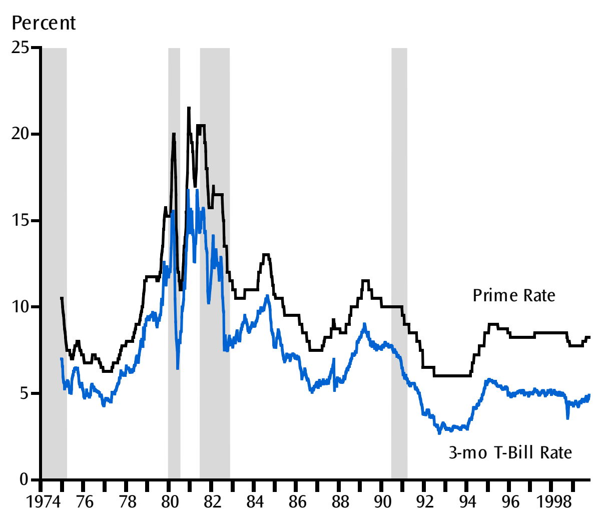
Table 1

Observation Categories Based on Size of Prime Rate Change. Sample period: 1/3/75 – 10/15/99

Category	*Criterion*	*Frequency*
1	ΔPrime $< -.50$	12
2	ΔPrime $= -.50$	44
3	ΔPrime $= -.25$	41
4	ΔPrime $= 0$	1105
5	ΔPrime $= +.25$	45
6	ΔPrime $= +.50$	33
7	ΔPrime $> +.50$	15

Figure 1

Prime and 3-Month T-Bill Rates



[1] If one were seeking to maximize a measure of fit for prime-rate changes, the target federal funds rate and/or the discount rate would be strong explanatory variables. Here, however, the objective is to examine how the prime rate moves relative to a benchmark short-term interest rate, the three-month T-bill rate.

to bring about increases and decreases in the prime rate are equal to each other:

$$\begin{aligned}
(5)\qquad \Delta Y_t &< -.50 \;\leftrightarrow\; Z_t < c_1 \\
\Delta Y_t &= -.50 \;\leftrightarrow\; c_1 \le Z_t < c_2 \\
\Delta Y_t &= -.25 \;\leftrightarrow\; c_2 \le Z_t \le c_3 \\
\Delta Y_t &= 0.00 \;\leftrightarrow\; c_3 < Z_t < c_4 \\
\Delta Y_t &= +.25 \;\leftrightarrow\; c_4 \le Z_t \le c_5 \\
\Delta Y_t &= +.50 \;\leftrightarrow\; c_5 < Z_t \le c_6 \\
\Delta Y_t &> +.50 \;\leftrightarrow\; Z_t > c_6.
\end{aligned}$$

The estimation results in Table 2 represent a run of the Gibbs sampler of 8,000 iterations, where the last 5,000 were saved to calculate posterior moments after the sampler had converged. Table 2 presents two sets of results, with and without the 1994-99 period. The results in Dueker (1999) showed that inclusion of Markov switching in the variance was necessary for a model of prime-rate changes to pass a chi-square goodness-of-fit test. The model presented here easily passes the same test with probability values of 0.50 for the 1975-99 period and 0.51 for the 1975-94 period. Therefore, I present the tests for asymmetry based on this specification of the model.

The estimated β coefficients in Table 2 show evidence of asymmetry in the form of sluggish downward response to decreases in the T-bill rate. As one would expect, the overall response (β_1) is not significantly different from one, for either increases or decreases. This means that the prime rate and the T-bill rates do not wander apart from each other in the long run. Moreover, it does not take long for the latent prime rate, Y^*, to make a one-to-one adjustment to a change in the T-bill rate—about eight weeks for decreases according to the estimates in Table 2.

For increases, on the other hand, the adjustment is even faster, as the lag coefficients beyond three weeks are not jointly significant (β_4). Thus, the latent prime rate, Y^*, makes a one-to-one adjustment with the T-bill rate within three weeks when rates are increasing, but does not make the same adjustment until eight weeks have elapsed when rates are decreasing. The probability value of the Wald test for equal β coefficients at all eight lags easily rejects in favor of the hypothesis that the prime rate takes longer to make a full adjustment to a decrease than an increase in the T-bill rate. Thus, a statistically significant asymmetry exists in the slope coefficients between increases and decreases; by itself, however, it is not of great economic importance whether full adjustment to the latent prime takes place in eight weeks or three weeks. For this reason, we also examine the contribution of the threshold coefficients to asymmetry in the prime rate.

Table 2 also contains estimates of the threshold coefficients, which can serve as additional source of asymmetry if we can reject $c_3 = -c_4$, $c_2 = -c_5$, and $c_1 = -c_6$. These threshold coefficients measure the impetus for a change in the prime that is needed to induce a discrete change. The two sets of results in Table 2 both show that the threshold coefficients are asymmetric between increases and decreases. The impetus $Z_t = Y_t^* - Y_{t-1}$ needed to change in the prime rate is substantially larger for a decrease of any size than for a corresponding increase. For a 25 basis-point decrease, the impetus must be greater than 90 basis points (c_3) according to the 1975-99 estimates vs. 44 basis points for a 25 basis-point increase in the prime. The fact that both c_3 and c_4 are greater than the 25 basis-point change that they demarcate demonstrates that the prime rate is fairly rigid. Across all three size categories, the magnitude of the asymmetry is between 40 and 50 basis points, and we can easily reject $c_3 = -c_4$, $c_2 = -c_5$, and $c_1 = -c_6$ at the 95-percent confidence level. For the data sample that ends in April 1994, the magnitude of the asymmetry is somewhat less, between 30 and 35 basis points, but is still highly statistically significant.

For either sample period, the impetus needed to induce a decrease in the prime rate is large enough to lift the average level of the prime rate relative to the mean of the latent prime, Y^*. The difference in their means is 25 basis points, with the average difference even greater in periods of falling interest rates. Figure 2 plots the latent prime Y^* and actual prime Y for roughly the last 400 weeks in the sample, which include periods of rising and falling interest rates. The figure shows that, during periods of rising interest rates, the actual prime rate does not have a mean appreciably different from that of the latent prime. Because of asymmetries in the threshold coefficients, however, the prime lags behind the latent prime when the two rates are falling. This asymmetry in the threshold coefficients leads to the gaps between the latent prime Y^* and actual prime Y shown in Figure 2 during periods of falling interest rates.

Table 2

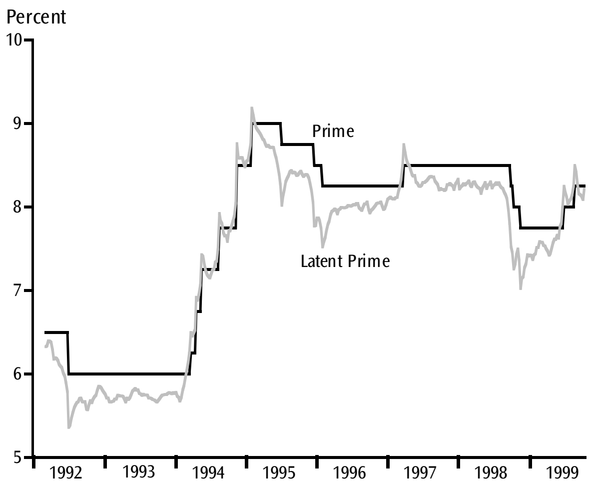
Posterior Distributions of Parameters

Parameter	1975-99 Post. mean	1975-99 95% interval	1975-94 Post. mean	1975-94 95% interval
Coefficients for changes in T-bill rate INC for increases: DEC for decreases				
β_1 **INC**	1.09	(.912,1.26)	1.09	(.911,1.26)
β_1 **DEC**	1.11	(.971,1.26)	1.12	(.966,1.26)
β_2 **INC**	.447	(.205,.695)	.434	(.145,.677)
β_2 **DEC**	.613	(.437,.790)	.616	(.436,.798)
β_3 **INC**	.327	(.099,.562)	.335	(.097,.571)
β_3 **DEC**	.492	(.310,.670)	.489	(.298,.662)
β_4 **INC**	−.020	(−.261,.236)	−0.46	(−.304,.196)
β_4 **DEC**	.408	(.223,.596)	.413	(.210,.618)
β_5 **INC**	−.041	(−.264,.189)	−.062	(−.283,.166)
β_5 **DEC**	.350	(.159,.539)	.355	(.163,.563)
β_6 **INC**	.010	(−.198,.226)	−.011	(−.220,.199)
β_6 **DEC**	.239	(.061,.413)	.244	(.056,.428)
β_7 **INC**	−.021	(−.211,.173)	−.051	(−.244,.133)
β_7 **DEC**	.182	(.031,.346)	.213	(.034,.398)
β_8 **INC**	.044	(−.124,.221)	.037	(−.139,.200)
β_8 **DEC**	.219	(.067,.368)	.242	(.074,.403)
Cut-off coefficients				
c_1	−1.44	(−1.64,−1.26)	−1.35	(−1.55,−1.21)
c_2	−1.05	(−1.22,−.926)	−.950	(−1.08,−.881)
c_3	−.929	(−1.09,−.812)	−.840	(−.963,−.773)
c_4	+.441	(.321,.571)	+.492	(.433,.559)
c_5	+.583	(.454,.712)	+.647	(.583,.712)
c_6	+.923	(.719,1.17)	+1.02	(.876,1.20)
$c_3 + c_4$	−.487	(−.760,−.252)	−.348	(−.505,−.264)
$c_2 + c_5$	−.467	(−.737,−.226)	−.304	(−.454,−.207)
$c_1 + c_6$	−.514	(−.867,−.156)	−.324	(−.569,−.105)
Markov switching drift coefficients				
β_0 $(S2 = 0)$	−.009	(−.037,.014)	−.010	(−.042,.016)
β_0 $(S2 = 1)$	.015	(−.008,.043)	.018	(−.008,.052)

In sum, we find evidence of two types of asymmetry in the prime rate. First, the latent prime rate moves more quickly in relation to the T-bill rate if the two rates are rising than if they are falling, but this time lag by itself is not likely to be of economic importance. Second, the threshold coefficients indicate that a larger impetus is required to lower the prime rate than to raise it. But at the same time, the thresholds imply that the prime rate will respond to sufficiently large gaps between the latent prime and the actual prime, so the gap cannot become arbitrarily large. Once a threshold is reached,

Figure 2

Asymmetric Behavior of Prime Rate in Relation to Latent Prime Rate



the prime will change to begin to close the gap. A gap can persist following a series of interest rate cuts, however, until interest rates turn upward, as shown in Figure 2.

SUMMARY AND CONCLUSIONS

This article attempts to move from anecdotal evidence about asymmetric behavior of bank loan rates to specific econometric tests. I use these tests to address the arguments that asymmetry in bank-administered interest rates implies that asymmetry renders borrowers more vulnerable to business cycle downturns than they otherwise would be and that monetary policy easings have less effect than monetary policy tightenings. Within the asymmetric bounds implied by the threshold coefficients, both of these arguments appear to hold, but it is important to note the limited scope of the asymmetry, which does not permit the prime to remain too far above where it would be in the absence of asymmetry. Similarly, it is perhaps too easy to jump to the conclusion that, other things being equal, borrowers would benefit if the asymmetry were not present. The asymmetry is probably the market's response to the rise in default and late payment probabilities of borrowers that occurs during cyclical downturns. Hence, fewer funds would be loaned in the absence of asymmetry, so it is not possible to remove the asymmetry and keep other things equal.

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