Market Structure and the Nature of Price Rigidity: Evidence from the Market for Consumer Deposits

Author(s): David Neumark and Steven A. Sharpe

Source: *The Quarterly Journal of Economics*, May, 1992, Vol. 107, No. 2 (May, 1992), pp. 657-680

Published by: Oxford University Press

Stable URL: https://www.jstor.org/stable/2118485

MARKET STRUCTURE AND THE NATURE OF PRICE RIGIDITY: EVIDENCE FROM THE MARKET FOR CONSUMER DEPOSITS*

DAVID NEUMARK AND STEVEN A. SHARPE

Panel data on consumer bank deposit interest rates reveal asymmetric impacts of market concentration on the dynamic adjustment of prices to shocks. Banks in concentrated markets are slower to raise interest rates on deposits in response to rising market interest rates, but are faster to reduce them in response to declining market interest rates. Thus, banks with market power skim off surplus on movements in both directions. Since deposit interest rates are inversely related to the price charged by banks for deposits, the results suggest that downward price rigidity and upward price flexibility are a consequence of market concentration.

I. INTRODUCTION

The issue of how prices respond to fluctuations in money supply is viewed by many macroeconomists as important, if not central, to understanding the relationship between fluctuations in the money stock and aggregate output. There still exists, of course, a great deal of disagreement over the appropriate interpretation of the observed correlation between money (or its counterpart, credit) and output. Following the Great Depression, and throughout the 1950s, a popular hypothesis ascribed this correlation to downward price rigidity, which was thought to be the result of inflexible, noncompetitive corporate pricing policies.[1] Beginning in the late 1970s, it became more popular to model the money-price-output relationship as largely driven by expectations. This was reflected in both "Keynesian" models as well as "equilibrium" models driven by intertemporal substitution.

In the last decade, though, macroeconomists have displayed a renewed interest in the nature of competition and pricing in product markets. This revival may be attributed in some part to the

*We would like to thank Swati Bhatt, John Duca, Cara Lown, Glenn Rudebusch, Peter Zemsky, and participants in the Federal Reserve Board research workshop and the Wharton School macrofinance seminar for helpful comments. Extensive research assistance was provided by Gerhard Fries. This research was completed while both authors were employed at the Board of Governors of the Federal Reserve System. The views expressed herein are solely those of the authors.

1. The hypothesis that connected market structure or corporate practices to price inflexibility, often credited to Means [1935], became known as the administered price hypothesis. See Scherer [1980] for a review and assessment of the "scores of quantitative studies" that tested the administered price hypothesis using price indices published by the Bureau of Labor Statistics.

The Quarterly Journal of Economics, May 1992

writings of Okun [1981], who emphasized why the special features of customer markets should lead to price rigidities. It may also be attributed to the rich array of pricing practices and patterns of market behavior advanced in the growing theory of industrial organization.

While most recent research linking market characteristics to price rigidity has been theoretical, there have been a few empirical studies that draw on microeconomic price data from various industries.[2] Because the key market or product characteristics stressed by many of the newer theories (such as information structure) are difficult to measure or observe, existing empirical studies have focused on that old workhorse: market concentration.

The purpose of this paper is to extend this line of empirical research by analyzing the impact of market concentration on the adjustment of prices in the market for consumer bank deposits. The placement of funds in deposits can be thought of as a purchase of investment and possibly transaction services, the price of those services being positively related to the spread between comparable open market interest rates and the interest rate paid on deposits. Since deposits are not a standard consumption good, the pricing of bank deposits per se may be thought to have no direct implications for the money-output linkage in the traditional sense outlined above. However, unless the effects of concentration in this market are thought to be entirely idiosyncratic, there is no reason why behavioral implications cannot be generalized to other markets.

As a study of the dynamic price-concentration relationship, this paper has two important advantages compared with previous research. First, while the price-concentration relationship is typically identified from variation *across industries,* this study focuses on a single industry, and identifies the impact of market concentration from variation *across geographical markets.* It thereby avoids problems posed by industry-specific factors such as diverse cost structure dynamics or varying union strength.[3] Here, most of the variation in marginal opportunity cost is the consequence of shocks to the "global" security market interest rate, shocks that are thus

2. Stigler and Kindahl [1970], well-known for pointing out the potential weaknesses of the early research that relied on published price indices, were the first to investigate these issues using actual transaction price data collected for a wide variety of commodities. A review of recent research employing micro panel data to analyze the determinants and extent of price rigidity can be found in Carlton [1989].

3. Domowitz, Hubbard, and Petersen [1986a, 1986b] attempt to get around this problem by analyzing average price-cost margins. They find an important role for unions in accounting for the relationship between concentration and price dynamics [1986b].

common to all geographical markets. Second, the high frequency of industrywide shocks yields a set of observations and price movements rich enough to uncover possible asymmetries in the effects of concentration; that is, any difference in the way concentration impacts downward versus upward price adjustment.[4]

Finally, there is an additional reason for studying the impact of market structure on price adjustment in this market. Since the deregulation of deposits in the early 1980s, the dynamics of deposit prices have been held responsible, at least in part, for a deterioration in the target and indicator properties of actual money supply measures. As pointed out by Simpson [1984], since deregulation money demand has been quite sensitive to fluctuations in the spread between open market interest rates and own rates on deposits. In the face of large swings in interest rates, sluggish deposit rate movements have given rise to temporary swings in money demand that have nothing to do with spending plans.[5] The current environment of deregulation, merger, and consolidation of the banking industry is likely to have a profound effect on industry structure. Since the dynamics of deposit rate adjustment have important implications for the functioning of monetary policy, it would be useful to understand how industry structure affects these dynamics.

In the following section, we lay out the conceptual framework for analyzing bank deposits and the determination of prices. In Section III we describe the data and report descriptive statistics and estimates of the partial effect of concentration under a static equilibrium characterization. In Section IV we estimate a partial adjustment model, modified to allow local market characteristics to affect speeds of adjustment. In Section V we develop and estimate a switching model that allows these effects to show directional asymmetry. In Section VI we summarize the results and draw implications.

The most interesting finding is the impact of market concentration on the adjustment of prices to shocks, and the asymmetric

4. Scherer [1980] concludes that, all told, studies using BLS industry price indices provide weak evidence of an asymmetric effect of concentration on price adjustment: more concentrated industries seem to display downward price rigidity but upward price flexibility. Carlton [1986] finds a positive, significant impact of concentration on the duration of price fixity, but no asymmetry in the pattern of that rigidity. He conjectures such findings may reflect the use of devices other than price for short-run allocation by large firms.

5. See Diebold and Sharpe [1990] for a thorough statistical account of market interest rate and deposit rate dynamics from 1983 through 1985, and Roth [1987], for example, for a recent analysis of the impact of deregulation on the functioning of monetary policy.

nature of the relationship between market structure and price adjustment. Banks in concentrated markets tend to be *slower* to *increase* interest rates on deposits in response to rising open market rates; on the other hand, these same firms are *faster* to *lower* interest rates on deposits in response to falling market interest rates. Thus, banks with market power skim off extra surplus on movements in both directions. Since the interest rate is the inverse of the price charged by the bank for deposits, this suggests more generally that downward price rigidity and upward price flexibility are the consequence of market concentration.

II. Conceptual Framework

Our conceptual model, like those of Flannery [1982] and Diamond [1984], for example, focuses on the bank's role as investment agent for households. Banks take savings in the form of deposits from households and lend these funds out for investment. Banks may be more efficient investors due to economies of scale, diversification benefits, or a reduction in total transaction and monitoring costs. In any case, if banks were perfectly competitive with respect to household depositors, they would pay a rate of return on deposits equal to the rate of return they earn on loans, less any costs of doing business, such as the cost of transaction services provided to depositors.

Due to banking regulations, the markets from which banks draw their deposits have generally been confined to small geographical areas, such as counties or metropolitan areas, or, in some cases, states.[6] The investment of bank funds, though, is done on a national scale, facilitated by the federal funds and treasury securities markets. Thus, through the interbank purchase and issuance of securities, the marginal cost of capital is equalized across local markets. Profit maximization by banks in their loan markets will thus result in the equating of this global marginal cost with the rate of return generated on the marginal investment or loan by each bank.[7]

6. Even in those states that allow statewide banking, many banks draw a substantial proportion of their deposits from one location, often the location of their home office. Gilbert [1984] summarizes evidence in favor of this market definition.

7. If banks are competitive in loan markets, then all borrowers also pay the (risk-adjusted) rate of return earned on the marginal loan (the global marginal cost of capital). If banks have some local monopoly power in loan markets, borrowers would in general pay a rate of return higher than the marginal cost of funds. As in Flannery [1982], our analysis and inferences concerning deposit market behavior are independent of the scenario that characterizes loan markets. This is owing to

Three departures from this perfectly competitive framework provide the basis for our empirical model. First, if banks in more concentrated deposit markets extract more of the surplus from the business of investing deposits, one would expect banks in concentrated markets to pay lower deposit interest rates. In fact, fairly strong empirical evidence of a negative cross-sectional relationship between market concentration and deposit interest rates is provided, for example, by Berger and Hannan [1989].[8]

Second, even the existence of competitive pricing does not necessarily imply that households earn the marginal rate of return at every moment in time. If the bank-depositor relationship is sufficiently durable over time, then competitive pricing implies only that deposits earn the marginal rate of return averaged over some horizon. Indeed, intertemporal variability in the spreads between aggregate deposit interest rates and security market interest rates is well documented.[9] There is no widely accepted explanation for the lack of deposit yield indexation to open market yields (or the lack of indexing in other parts of the economy, for that matter). Nonetheless, given the absence of indexation, it is perhaps not surprising that deposit yields are not updated with every movement of the T-Bill or federal funds rate. While "menu costs" of repricing may be relatively low compared with repricing in other markets, they are not necessarily trivial. On the other hand, as suggested by Flannery [1982], banks may find it costly to induce rapid adjustments in their stock of deposits, and it is costly for consumers to frequently search out the best deal and move funds among banks every time a few basis points could be gained. Consequently, both consumers and banks may benefit from a pricing practice that merely reflects the market yield on average.

Finally, it is plausible that deposit price sluggishness varies systematically across banks. In particular, the dynamics of deposit

---

the presumption of a competitive interbank funds market; isomorphic to the classic separation of saving and investment decisions under perfect capital markets, here, the marginal revenue any bank earns on its marginal loan is equated with its marginal cost of funds, or the interbank security rate.

8. They uncover a persistent and significant negative correlation which is strongest when interest rates are at their peak within their 1983–1985 sample period. Cross-sectional studies of deposit pricing prior to the elimination of regulation Q, reviewed in Gilbert [1984], provide weaker evidence of a negative concentration effect on yields. None of these studies examine the relationship between concentration and dynamics.

9. Flannery documents episodes of pre-deregulation deposit interest sluggishness. Moore, Porter, and Small [1988] empirically model the relationship between the federal funds rate and the slower-moving aggregate deposit rates in the form of a cointegration model. Diebold and Sharpe [1990] perform a multivariate analysis of post-deregulation deposit and market interest rate dynamics.

prices may be related to market structure. For example, along the lines of the administered price hypothesis of Means [1935] or the kinked demand curve theory of Hall and Hitch [1939], concentrated markets may display greater deposit interest rate rigidity.[10]

Reflecting these considerations, a dynamic model of deposit pricing, competitive or otherwise, may be represented quite generally by the discrete pricing function:

$$y_{it} = \psi[\phi(x_{it},R_t),x_{it},y_{i,t-1},R^e_{i,t+1}], \tag{1}$$

where y_{it} is the yield on bank i's deposit account at time t. R_t is the (global) marginal cost of capital, or equivalently, the yield on the intermarket loan rate of the appropriate maturity, x_{it} is a vector of bank- or market-specific characteristics, and $R^e_{i,t+1}$ is the bank's subjective probability distribution over the path of future marginal yields. $\phi(\cdot)$ represents the "steady state" equilibrium pricing relationship, and $\psi(\cdot)$ represents the actual dynamic behavior. As suggested in equation (1), x_{it} may affect the dynamics of prices as well as the division of surplus in steady state equilibrium.

III. Data and Preliminary Results

This study focuses on the yield behavior of two types of retail deposits: the six-month certificate of deposit (6MCD) and the money market deposit account (MMDA). The first is a time deposit, in which funds are often left to be rolled over into a new CD upon maturation. The second is a savings/limited transaction account with no term to maturity. The interest rates offered on these accounts were reported monthly by approximately 400 banks in a survey conducted by the Federal Reserve Board. The yields reported by each bank correspond to those most commonly paid on the account in question during the week prior to the final Wednesday of the month.

The final sample used is a panel of 255 banks located in 105 markets over a period of time beginning in October 1983 and ending in November 1987.[11] Only those banks located in Metro-

10. In a related study of deposit rate dynamics, Hannan and Berger [1991] explore the testable implications of Rotemberg and Saloner's [1987] version of the menu cost theory of price rigidity.

11. On October 1, 1983, rate ceilings were removed on all time deposits. Furthermore, MMDAs were only first introduced by some institutions in January of 1983. Many rates offered on these accounts throughout the summer of 1983 were inflated introductory offers. Because of this, and for purposes of convenience, we chose to begin both series in October.

politan Statistical Areas (MSAs), with at least two thirds of their deposits drawn from branches in their "home" market, are included. This set of restrictions eliminates what are perhaps the least sophisticated banks, located in outlying counties. It also eliminates one- and two-bank towns, where households probably have relatively few alternatives. In such markets there is greater potential for interdependence between loan market and deposit market services and prices, and thus, some distortion in our measure of the rate of return earned on deposits. Finally, the "home market" restriction is used in order to assure the meaningfulness of the measures of local market concentration.

The measure of market concentration employed is the Herfindahl index, computed using FDIC data on total deposits of each bank and S&L located in any given market.[12] The Herfindahl index is simply computed as the sum of squared market shares of total bank and S&L deposits in the market. It thus ranges between zero, for atomistic competition, to one, for monopoly. In standard goods market models, if firms are static Cournot competitors facing constant marginal costs, then the price-cost margin can be shown to be proportional to the Herfindahl index.

As in previous studies, we also include controls for variation in branching restrictions across states. All else equal, one might expect branching restrictions to result in lower deposit yields on average, due to the greater barriers to entry they pose. They may also increase costs in smaller markets by making it difficult for banks to exhaust economies of scale. LIMIT is a dummy variable which takes on the value of one when only limited (statewide) branching is allowed. UNIT takes on a value of one if no branching is allowed.

A measure of yearly population growth by SMSA, obtained from a Bureau of Labor Statistics survey, was also included. This variable (POPGRO) proxies the percentage of new customers in the market. Because new households in an area may be more prone to searching for the best yield opportunities than previous market inhabitants, one might expect high growth areas to be more competitive, all else equal.

Finally, our proxy for the marginal cost of funds is the yield to

12. We are grateful to Berger and Hannan for providing these indices, which were constructed for their [Berger and Hannan, 1989] cross-sectional study. These indices were computed yearly through 1985. Due to problems with FDIC data in the later years, we did not extend this series. Since variation in the index over time was very small in comparison with cross-sectional variation, using the 1985 measures for 1986–1987 does not pose a problem.

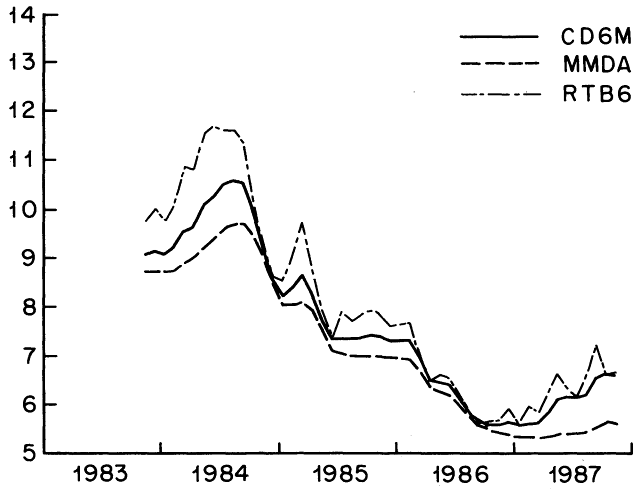

FIGURE I

maturity on the six-month treasury bill, reported on the Wednesday (each month) one week prior to the date when banks reported their deposit rates. This day corresponds to the beginning of the week for which the weekly average deposit rates were computed by each bank.[13] In Figure I the time series of the treasury bill is plotted along with the average yield offered on MMDAs and 6MCDs for our sample of banks. The relative sluggishness of average sample deposit rates is quite clear.

Summary statistics describing the variation of market concentration and population growth across geographical markets, as well as pairwise correlations among all variables employed, are given in Table I. All statistics are computed using individual bank means, computed over the time span of the data set.[14] For reference purposes the distribution of the Herfindahl index over banks

13. The T-Bill rate was chosen over the federal funds rate because of its more appropriate maturity. Alternatively, one might use the wholesale, or "jumbo" CD rate, the interest rate on large CDs which trade in a competitive secondary market. VAR-style time series analysis on various aggregate deposit and market interest rates performed by Diebold and Sharpe [1990] found the time series behavior of the six-month wholesale CD rate to be nearly indistinguishable from that of the six-month T-Bill rate.

14. Because most of the variation in market structure comes from cross-sectional observations, the numbers in this table are nearly identical when computed for observations, rather than bank means. Furthermore, since most variation in deposit rates is explained by market interest rate movements over time, using bank means to compute correlations gives a much better reading of the role of cross-sectional influences.

TABLE I
SUMMARY STATISTICS: BANK MEANS*

	Herfindahl index (HERF)	Three-firm conc. (CR3)	No. of firms (NUMBER)	Population growth (POPGRO)
Mean	0.083	0.402	109	0.008
Std. dv.	0.040	0.114	103	0.010
Minimum	0.031	0.186	9	−0.022
Maximum	0.256	0.770	435	0.053

Pearson correlation coefficients

	CR3	NUMBER	POPGRO	LIMIT	UNIT	6MCD	MMDA
HERF	0.97	−0.47	−0.06	0.36	−0.21	−0.41	−0.44
CR3		−0.41	−0.02	0.39	−0.20	−0.37	−0.43
NUMBER			0.04	−0.22	0.50	0.15	0.22
POPGRO				−0.19	0.24	0.24	0.05
LIMIT					−0.35	−0.27	−0.33
UNIT						0.11	0.15
6MCD							0.55

*All statistics are computed using the banks' mean values for each variable. The sample comprises 49 monthly observations on each of 255 banks, a total of 12,495 observations.

(computed using both bank and S&L market deposits) is supplemented by that of the three-firm concentration ratio (CR3) and the number of firms per market.

There is quite a wide range of concentration. The CR3 ranges from 19 percent to 77 percent, although most banks reside in markets in the lower half of this range. The average CR3 is 40 percent, and its standard deviation is 11 percent. Correspondingly, the average HERF is 0.083, while its standard deviation is 0.040, and it ranges between 0.031 (in Nassau-Suffolk, NY) and 0.256 (in Laredo, TX). Finally, the smallest number of firms is 9 (also in Laredo), and the largest number is 435 (in the city of New York).

In the bottom half of the table, correlations are given for pairwise relationships among the variables. Most notable is the substantial negative correlation between the deposit rates and the Herfindahl index: −0.41 for 6MCD, and −0.44 for MMDA. Finally, while not reported in the table, it is worth noting that about one fifth of our observations are in markets with unit banking, and about one third are from markets with limited branching. The only change over time is in 1987, when several states moved toward fewer restrictions.

In Table II we report the results from regressions of bank

TABLE II
BANK MEAN REGRESSIONS: OLS*

	6MCD, long (1)	6MCD, short (2)	MMDA, long (3)	MMDA, short (4)
HERF	−2.56	−2.64	−2.79	−2.66
	(0.42)	(0.43)	(0.46)	(0.46)
POPGRO	5.35	5.69	−0.37	−0.11
	(1.46)	(1.45)	(1.61)	(1.55)
LIMIT	−0.06	−0.06	−0.14	−0.10
	(0.04)	(0.04)	(0.04)	(0.04)
UNIT	−0.03	−0.03	0.01	0.05
	(0.04)	(0.05)	(0.05)	(0.05)
$\bar{R}^2$	0.21	0.21	0.21	0.18
Sample size	12,495	9,690	12,495	9,690
σ^2	0.06	0.06	0.07	0.07

*Regressions of bank mean deposit rate on bank mean values of explanatory variables. Standard errors are given in parentheses.

mean deposit rates on bank means of all four explanatory variables. These regressions might be thought of as "equilibrium" regressions; that is, they provide estimates of the structural effect of the explanatory variables under the assumption that markets are always in "equilibrium," in the static sense.

The first regression reported for each rate was run with means computed using the entire sample time period; the second was run with means computed over a shorter sample period that excluded the first year of observations. The shorter period was considered for two reasons. First, the early part of the sample may have been characterized by learning on the part of banks, since key restrictions were lifted only shortly before our sample period begins. Second, as can be seen in Figure I, the earlier part of the sample was characterized by sharply rising and then falling interest rates, while the later years experienced smaller fluctuations. The regressions on the short samples are robustness checks with respect to these potentially important factors.

In all four regressions the coefficient on the Herfindahl index is nearly identical, negative, and highly significant. The coefficients of −2.6 to −2.8 imply that a one-standard-deviation increase in HERF (0.04) is associated, on average, with a decline of about ten to eleven basis points in the bank mean deposit rate. Finally, the change in HERF associated with a move from the least to most concentrated market implies a drop of about 60 basis points in both deposit rates, on average.[15]

IV. The Partial Adjustment Model

In order to model the dynamics of deposit interest rates in an environment with fluctuating marginal costs of capital, a functional form is needed for (1). We employ an adaptation of the Koyck distributed lag scheme, also known as the partial adjustment model. Letting $y^*_{it} = \phi(x_{it}, R_t)$ represent the (long-run) equilibrium relationship, the Koyck scheme can be written in the following form:

$$y_{it} - y_{i,t-1} = \lambda(x_{it}) \cdot [\phi(x_{it}, R_t) - y_{i,t-1}] + \epsilon_{it}. \tag{2}$$

Our interpretation of (2) is nonstructural, in contrast to recent Euler equation derivations of partial adjustment, where y^* is derived from individual maximization subject to quadratic adjustment costs. Here, the model is more heuristic, a parsimonious method of approximating dynamics while at the same time estimating low-frequency, or "long-run," relationships. Since we are interested in the effects of market characteristics on adjustment speeds, as well as long-run pricing, the empirical model allows the "speed of adjustment," λ, to depend upon these market characteristics.

To estimate the model, we assume ϵ_{it} to be i.i.d. normal and make the functional form assumptions (3) and (4):

$$\phi(x_{it}, R_t) = R_t \cdot (x_{it}\beta_x) \tag{3}$$

$$\lambda(x_{it}) = x_{it}\lambda_x, \tag{4}$$

where the row vector of bank characteristics, x_{it}, has a one in the first column. Long-run equilibrium deposit rates are assumed to be proportional to the T-Bill rate, which is the secondary market yield to maturity on the six-month T-Bill on the second-to-last Wednesday each month (the first day of the week for which banks report week-average deposit interest rates). The factor of proportionality, as well as the speed of adjustment, is linear in the market characteristics.

15. These results are, not surprisingly, quite close to those of Berger and Hannan [1989] in their pooled regression. Each regression was also run with the inclusion of a variable equal to the total number of firms. In no case did this variable come in significantly or result in a higher adjusted R^2.

It should be emphasized that a finding that the fraction of adjustment is substantially less than one for the *average* bank should be interpreted with care. All that a low estimate of $x_{it}\lambda_x$ necessarily implies is that a bank's current period deposit rate has a relatively high correlation with its previous period deposit rate, compared with its correlation with the current month's T-Bill rate. This may reflect expectations of mean reverting market interest rates, costs of adjusting prices, or even the possibility that the six-month T-Bill is a noisy measure of marginal opportunity cost.[16]

However, such factors are much less likely to underlie cross-sectional differences in adjustment speeds associated with market structure. The estimates of the coefficients measuring the impact of market structure on the equilibrium markups and "speeds of adjustment" are thus potentially more informative. A negative estimate of λ_{HERF}, for example, would imply that the correlation of deposit rates with the T-Bill rate, relative to their correlation with the previous month's deposit rate, tends to be lower for banks in more concentrated markets. Such a finding would clearly constitute evidence for slower adjustment of prices by firms in oligopolistic markets—the prediction of the administered price hypothesis.

In Table III we report maximum likelihood estimates of the partial adjustment model (2)–(4), where the vector x includes a constant and the market characteristics, HERF, POPGRO, LIMIT, and UNIT. We again provide estimates for both the full and the shorter time series of 6MCD and MMDA rates.[17]

Estimates of the average markup, $\bar{x}\hat{\beta}$ (or, actually, "markdown" in this case), are the same in the short and long samples, about 0.93 for 6MCD and 0.86 for MMDA. Estimates of the average speed, or percentage, of adjustment per month, $\hat{\lambda}\,\bar{x}$, are approximately equal to

16. Theoretically, time aggregation could also pose a problem, since banks report a weekly deposit rate. In practice, this subtlety is probably unimportant. First, banks rarely change deposit rates more than once a week, and second, the time series consists of observations that are four weeks apart.

17. The estimates reported are computed from samples that excluded 32 outliers. Outliers were defined as observations for which estimated residuals from partial adjustment model specifications were greater than 1.5 in absolute value (over six times the standard errors of these specifications when the entire sample is used). Without dropping a subset of these outliers, convergence problems were encountered in estimating the most general asymmetric models that follow. The estimates presented in Table III (and the statistics in Tables I and II) were insensitive to dropping these observations. The estimates of those asymmetric models that did converge using the entire sample were qualitatively unaffected by the omission of outliers. If a bias is created by dropping these outliers, it is most likely to be a bias toward the acceptance of the symmetry restriction that is imposed in the model used to estimate these residuals, since these were outliers *least* consistent with the partial adjustment model.

TABLE III
PARTIAL ADJUSTMENT MODEL: MAXIMUM LIKELIHOOD*

	6MCD, long (1)	6MCD, short (2)	MMDA, long (3)	MMDA, short (4)
β_0	0.955	0.964	0.904	0.897
	(0.002)	(0.003)	(0.003)	(0.004)
λ_0	0.333	0.352	0.253	0.252
	(0.007)	(0.010)	(0.005)	(0.007)
β_{HERF}	−0.312	−0.378	−0.412	−0.384
	(0.021)	(0.032)	(0.029)	(0.042)
λ_{HERF}	−0.021	−0.085	−0.173	−0.062
	(0.080)	(0.113)	(0.050)	(0.069)
β_{POPGRO}	0.698	0.805	−0.233	−0.540
	(0.066)	(0.078)	(0.093)	(0.136)
λ_{POPGRO}	0.185	0.620	−1.006	−1.418
	(0.245)	(0.340)	(0.151)	(0.195)
β_{LIMIT}	−0.007	−0.002	−0.017	−0.017
	(0.002)	(0.003)	(0.003)	(0.004)
λ_{LIMIT}	0.043	0.073	−0.013	−0.016
	(0.007)	(0.009)	(0.005)	(0.006)
β_{UNIT}	−0.007	−0.003	−0.002	0.014
	(0.002)	(0.003)	(0.003)	(0.004)
λ_{UNIT}	0.078	0.121	0.001	0.030
	(0.007)	(0.010)	(0.005)	(0.006)
σ^2	−0.070	0.069	0.044	0.040
	(0.001)	(0.001)	(0.000)	(0.000)
Sample size	12,463	9,670	12,463	9,670
Log likelihood	−1,084.1	−810.8	1,817.2	1,837.4

*Estimates shown are for sample excluding observations with absolute value of estimated residuals (from the partial adjustment model) greater than 1.5. Standard errors are given in parentheses.

the constant, $\hat{\lambda}_0$ in each regression—about 0.35 for both 6MCD regressions and 0.25 for both MMDA regressions. As in Diebold and Sharpe 1990, interest rates on MMDAs appear more sluggish than those on CDs.

The effect of market concentration on the long-run equilibrium markup is negative and highly significant in all four regressions, slightly more so for MMDA than for 6MCD. The average estimate of β_{HERF} of about −0.37 implies that a one-standard-deviation (0.04) increase in HERF results in a drop in the proportional markup by 0.014; for the average sample period T-Bill rate of about 7.5, this implies an eleven-basis-point drop in the deposit rate. The implied effect (again, given a T-Bill rate of 7.5) on equilibrium markups of going from the least to the most concentrated market is a decrease of about 60 basis points. Thus, our partial adjustment model estimates of the long-run equilibrium effect of concentration are very close to those estimates based upon the static bank means regressions reported in Table II.

On the other hand, estimates of λ_{HERF} suggest no clear effect of market concentration on the speed of adjustment. The coefficient is negative in all cases; but, in all but the full MMDA sample, it is far from significant. In the latter case it is significant, and its effect is small. A one-standard-deviation increase in HERF implies a decrease in the adjustment speed of about 0.007, less than 4 percent of the average adjustment speed of 0.25.

V. Partial Adjustment and Asymmetric Behavior

Implicit in the model above is the assumption that market adjustment toward long-run equilibrium is symmetric; that is, it does not depend upon whether deposit rates are above or below long-run equilibrium. The speed of adjustment, and the impact of market characteristics on that speed, are simply proportional to the degree of misalignment from the long-run relationship.

In an extensive study of the dynamics of money demand, however, Moore, Porter, and Small [1988] find evidence of asymmetry in the dynamic behavior of *aggregate* MMDA and NOW account rates. Deposit rates appear to adjust toward their supposed equilibrium level more promptly when they are above that level than when they are below. While this evidence is somewhat weak, owing to the paucity of observations, there is a striking similarity across the types of the deposit accounts they examine.[18]

With our panel data set, degrees of freedom should pose no obstacle. More importantly, if such asymmetry exists, it will be possible to test whether it is related to market structure. This could provide much needed clues in the search for a theory of asymmetric price rigidities.[19] While the theoretical microfoundations literature

18. Moore et al. estimate the rate dynamics in an error-correction framework, where the deposit rate is assumed to be linearly related to the federal funds rate in stationary equilibrium. Considerable structure was imposed on the estimation of the equilibrium relationship which was then used to estimate asymmetric dynamics. The authors conjecture that the asymmetry may reflect a response to differential behavior across investors, some of whom are yield maximizers, and others of whom focus on the convenience of customary transaction arrangements.

19. Theories generating price rigidity from menu costs or kinked demand curves generally rationalize only symmetric price rigidity. As noted in Rotemberg's [1987] discussion of the price rigidity literature, theories of kinked demand curves based upon the Nash equilibrium framework tend to give rise to a continuum of equilibria. Prices are "sticky" in the sense that the set of equilibria following a small

has had little to say about asymmetric stickiness, apparent asymmetries in price adjustment have been acknowledged in several contexts. Perhaps the most relevant of these is that the prime rate appears to be slower to decline in response to a drop in funding costs than it is to rise in response to an increase in funding costs. Arak et al. [1983] ascribe this finding to cyclicality in the elasticity of loan demand faced by banks.[20] Ausubel [1989] finds downward stickiness in credit card interest rates.[21] Finally, a tangential though perhaps not entirely coincidental observation is the "institutionalist" perspective of the labor market which ascribes asymmetric wage behavior (downward rigidity) to the (monopolistic) market power of unions.[22]

In order to explore the possibility of asymmetric behavior, we estimate a switching model of partial adjustment. This model allows adjustment speeds to vary according to whether a bank is above or below its long-run equilibrium markup, and includes the partial adjustment model (2)–(4) as a special case. It simultaneously estimates both the long-run relationship and the upward and downward adjustment speeds, all as functions of market factors.

The stochastic switching model of deposit rate adjustment is composed of three equations. (Time and bank subscripts are dropped for notational simplicity.) The first is a stochastic indicator of the regime—above or below long-run equilibrium—in which an observation lies. The other two are the stochastic equations of motion in each regime, in which the speeds of adjustment, $\lambda_{x(A)}$ (above equilibrium) and $\lambda_{x(B)}$ (below equilibrium), may differ.

shock does not change much, and thus is likely to include the preshock equilibrium. Such a structure is clearly symmetric as long as the shocks are distributed symmetrically.

20. They argue that the prime rate is determined by competition between banks for relatively small borrowers, and that loan demand elasticity of such borrowers is relatively low during recessions—times of falling interest rates—but high during expansions. The reason for this, they argue, is that small borrowers are hard-pressed to find alternative sources of finance, aside from their old bank, during recessions; this, in turn, may be the result of adverse selection problems lenders may face at such times. Forbes and Mayne [1989] present similar empirical evidence.

21. The stickiness is conjectured to be the result of both search costs and adverse selection, though the adverse selection arises due to a specific form of irrationality on the part of credit card users.

22. According to Bok and Dunlop [1970, p. 287], "Unions have had an effect on the secular rise of prices by contributing to the virtual disappearance of wage decreases. In the nineteenth century, long-term declines took place in money wages and prices in addition to short-run cuts in both during the depression years. . . . In the past generation, however, money wages and the price level have moved only upward."

Specifically,

$$I^* = y_{-1} - [R \cdot (x\beta_x) + \epsilon_E] \tag{5}$$

$$\Delta y_A = (x\lambda_{x(A)}) \cdot [(R \cdot (x\beta_x) + \epsilon_E) - y_{-1}] + \epsilon_A \tag{6}$$

$$\Delta y_B = (x\lambda_{x(B)}) \cdot [(R \cdot (x\beta_x) + \epsilon_E) - y_{-1}] + \epsilon_B. \tag{7}$$

We define composite errors in (6) and (7): $\tilde{\epsilon}_A \equiv (x\lambda_{x(A)}) \cdot \epsilon_E + \epsilon_A$ and $\tilde{\epsilon}_B \equiv (x\lambda_{x(B)}) \cdot \epsilon_E + \epsilon_B$; ϵ_E enters these errors because we now assume that $R \cdot (x\beta_x)$ measures the equilibrium rate with error. The "underlying" errors are assumed to be normal:[23]

$$(\epsilon_E,\epsilon_A,\epsilon_B) \sim N(0,\Sigma),\ \Sigma = \begin{bmatrix} \sigma_{EE} & 0 & 0 \\ 0 & \sigma_{AA} & 0 \\ 0 & 0 & \sigma_{BB} \end{bmatrix}.$$

The likelihood is thus

$$\begin{aligned} \Pr(I^* > 0,\tilde{\epsilon}_A) + \Pr(I^* \le 0,\tilde{\epsilon}_B) &= \Pr(I^* > 0|\tilde{\epsilon}_A)\Pr(\tilde{\epsilon}_A) \\ + \Pr(I^* \le 0|\tilde{\epsilon}_B)\Pr(\tilde{\epsilon}_B) &= \Pr(\epsilon_E < y_{-1} - R \cdot (x\beta_x)|\tilde{\epsilon}_A) \\ &\times \Pr(\tilde{\epsilon}_A) + \Pr(\epsilon_E \ge y_{-1} - R \cdot (x\beta_x)|\tilde{\epsilon}_B)\Pr(\tilde{\epsilon}_B). \end{aligned}$$

Under our assumption of joint normality of the structural errors,

$$E(\epsilon_E|\tilde{\epsilon}_A) = (\sigma_{E\tilde{A}}/\sigma_{\tilde{A}\tilde{A}}) \cdot \tilde{\epsilon}_A$$
$$\mathrm{var}(\epsilon_E|\tilde{\epsilon}_A) = \sigma_{EE} - \sigma^2_{E\tilde{A}}/\sigma_{\tilde{A}\tilde{A}},$$

and similarly for $\tilde{\epsilon}_B$. Thus, the likelihood can be written as

$$\Phi\left[\frac{(y_{-1} - R \cdot (x\beta_x)) - (\sigma_{E\tilde{A}}/\sigma_{\tilde{A}\tilde{A}}) \cdot \tilde{\epsilon}_A}{(\sigma_{EE} - \sigma_{E\tilde{A}}/\sigma_{\tilde{A}\tilde{A}})^{1/2}}\right] \cdot \phi(\tilde{\epsilon}_A) + \left(1 - \Phi\left[\frac{(y_{-1} - R \cdot (x\beta_x)) - (\sigma_{E\tilde{B}}/\sigma_{\tilde{B}\tilde{B}}) \cdot \tilde{\epsilon}_B}{(\sigma_{EE} - \sigma_{E\tilde{B}}/\sigma_{\tilde{B}\tilde{B}})^{1/2}}\right]\right) \cdot \phi(\tilde{\epsilon}_B), \tag{8}$$

where Φ is the standard normal distribution, and ϕ the normal density.[24] Given the definitions of $\tilde{\epsilon}_A$ and $\tilde{\epsilon}_B$, the variances and

23. σ_{AE} and σ_{BE} are assumed to be zero; however, the structural errors $\tilde{\epsilon}_A$ and $\tilde{\epsilon}_B$ are correlated with ϵ_E.

24. Unlike in most switching models (or even a simple probit), σ_{EE} is identified. Identification comes from two sources: y_{-1} appears in the arguments of the normal distribution function, with a coefficient constrained to equal one; and β_x appears alone in the arguments of the cumulative normal distribution term, and not just in the form β/σ_E, because β_x appears in both the indicator equation and the equations of motion.

covariances in the likelihood function are themselves functions of $\lambda_{x(A)}$ and $\lambda_{x(B)}$, as well as σ_{AA}, σ_{BB}, and σ_{EE}.

Estimation of the switching model is performed under the assumption that the underlying errors in adjustment, ϵ_A and ϵ_B, have identical variance: $\sigma_{AA} = \sigma_{BB} = \sigma^2$. In fact, without some such restriction on the relative size of these variances, the likelihood function above is unbounded [Maddala, 1983, p. 183].

Before estimating the fully stochastic switching model, we examine a restricted version of particular interest. Specifically, we estimate the model under the additional restriction that the error variance in the long-run equilibrium equation, the switching equation, is zero:

$$\sigma_{EE} = \sigma_{E\tilde{A}} = \sigma_{E\tilde{B}} = 0. \tag{9}$$

Consequently, $\tilde{\epsilon}_A = \epsilon_A$, and $\tilde{\epsilon}_B = \epsilon_B$. Under this restriction the numerator in the arguments of the cumulative distribution terms of the likelihood function is just $(y_{-1} - R \cdot (x\beta_x))$, and the denominator is zero. The cumulative distribution term is thus either one or zero (the upper limit of integration is either positive or negative infinity) depending upon whether $(y_{-1} - R \cdot (x\beta_x))$ is positive or negative. Consequently, each observation is assigned to either one regime or another.

We call this version of the model an "exact switching model." It is of interest because its stochastic structure only differs from that of our partial adjustment model, (2)–(4), in that the restriction of symmetry is lifted. Thus, we can form a likelihood ratio statistic that tests solely for the existence of asymmetry in adjustment behavior under the null hypothesis that all error is attributed to errors in adjustment.[25] In order to maximize the likelihood function under assumption (9), it is necessary to use a reparameterization that sidesteps the problem of division by zero. Letting $d \equiv y_{-1} - R \cdot (x\beta_x)$, we reparameterize the likelihood function of the exact switching model:

$$[(d + |d|)/2d]\phi(\epsilon_A) + [(d - |d|)/2d]\phi(\epsilon_B). \tag{10}$$

The first factor in the square brackets equals one if $d > 0$ (deposit rate is above equilibrium) and equals zero if $d < 0$ (below

25. In the single equation symmetric partial adjustment model, two separate sources of error cannot be distinguished. Our formulation, (2)–(4), implicitly attributes the error to errors in adjustment.

equilibrium). The opposite holds for the second term in square brackets.[26]

Maximum likelihood estimates are reported in Table IV. The first set of estimates for each of the four samples corresponds to the constrained, or exact switching, model (5)–(10); the second set (in columns denoted by a prime) corresponds to the fully stochastic switching model (5)–(8).

The statistics of greatest interest are contained in Table IV.[27] These include the implied means of the estimated equilibrium markups, $\bar{\beta}$, and upward and downward adjustment speeds, $\bar{\lambda}_A$ and $\bar{\lambda}_B$. Notably, in each case the implied mean speed of adjustment from above is substantially greater than that from below equilibrium, confirming the aggregate time series results of Moore et al. [1988]. The likelihood ratio tests for the symmetry restriction are computed from log likelihood values given in Table III (partial adjustment model) and the first column of each pair in Table IV (exact switching model). Symmetry is overwhelmingly rejected at conventional significance levels; the test statistic, distributed as χ^2 with five degrees of freedom, one for the constant and each of the four characteristics, exceeds 250 in all cases.

A comparison of the constrained (exact) and unconstrained (stochastic) model estimates reveals no clear qualitative differences in their implications. In all cases, the coefficient estimates indicate strong asymmetries in mean adjustment speeds ($\bar{\lambda}_A > \bar{\lambda}_B$). Nonetheless, a likelihood ratio test of the constraint that σ_{EE} is zero leads to strong rejection of the constrained model. This is also clear from the estimates of σ_{EE} and σ^2. As must be the case if σ_{EE} is greater than zero, the variance of the adjustment error, σ^2, is smaller in the unconstrained model. In all four cases it drops by 35 percent to 50 percent; thus, nearly half the unexplained variance is attributed to variance in long-run markups.

Perhaps most interesting is the fact that, in all eight switching regressions, $\lambda_{\mathrm{HERF}(B)}$ is negative, while $\lambda_{\mathrm{HERF}(A)}$ is positive. For MMDAs the differences in the effects of the Herfindahl index on the speed of adjustment are always statistically significant; estimates of $\lambda_{\mathrm{HERF}(B)}$ are several standard errors below zero, while those of

26. Rudebusch [1986, 1987] suggests a generic "exact" formulation such as this as a general substitute for stochastic switching models, which are often poorly identified even in simple linear models. The problem of computational complexity is amplified in our case because of the high degree of nonlinearity.

27. The full set of estimates is provided in the working paper version of this paper.

TABLE IV
ASYMMETRIC PARTIAL ADJUSTMENT MODEL: MAXIMUM LIKELIHOOD*

	6MCD, long		6MCD, short		MMDA, long		MMDA, short	
	(1)	(1′)	(2)	(2′)	(3)	(3′)	(4)	(4′)
$\overline{\beta}$	0.947	0.960	0.947	0.962	0.897	0.927	0.890	0.932
$\overline{\lambda}_A$	0.451	0.497	0.448	0.500	0.338	0.406	0.320	0.420
$\overline{\lambda}_B$	0.261	0.229	0.311	0.264	0.113	0.088	0.092	0.061
β_{HERF}	−0.264	−0.250	−0.304	−0.213	−0.141	−0.121	−0.123	−0.200
	(0.036)	(0.043)	(0.045)	(0.045)	(0.033)	(0.046)	(0.034)	(0.049)
$\lambda_{HERF(\mathrm{A})}$	0.346	0.182	0.267	0.398	0.866	0.997	0.815	0.741
	(0.177)	(0.204)	(0.192)	(0.202)	(0.114)	(0.170)	(0.105)	(0.172)
$\lambda_{HERF(\mathrm{B})}$	−0.348	−0.458	−0.617	−1.153	−0.484	−0.412	−0.460	−0.263
	(0.154)	(0.105)	(0.287)	(0.125)	(0.074)	(0.037)	(0.154)	(0.053)
σ_{EE}	. . .	0.160	. . .	0.205	. . .	0.265	. . .	0.296
		(0.009)		(0.013)		(0.007)		(0.008)
σ^2	0.068	0.046	0.069	0.035	0.042	0.027	0.038	0.019
	(0.001)	(0.001)	(0.001)	(0.002)	(0.000)	(0.000)	(0.000)	(0.000)
Sample size	12,463	12,463	9,670	9,670	12,463	12,463	9,670	9,670
Log likelihood	−953.7	−884.7	−770.1	−683.6	2,154.2	2,349.9	2,036.9	2,430.7

*$\overline{\beta}$, $\overline{\lambda}_A$, and $\overline{\lambda}_B$ are the equilibrium markup and adjustment speeds for a bank with the average characteristics. Estimates shown are for sample excluding observations with absolute value of estimated residuals (from the partial adjustment model) greater than 1.5. Standard errors are given in parentheses. The same control variables used in Table III were also included in the regressions, with independent effects on the equilibrium markup, and upward and downward speeds of adjustment.

$\lambda_{\mathrm{HERF}(A)}$ are several standard errors above zero. The evidence is somewhat weaker for CDs, since the estimate of $\lambda_{\mathrm{HERF}(A)}$ is significant (at the 5 percent level) in only two of the four cases. For CDs, the p-values of the likelihood ratio test of the restriction that $\lambda_{\mathrm{HERF}(B)} = \lambda_{\mathrm{HERF}(A)}$ are 0.05 for both the full and the short sample estimates of the constrained model, and 0.11 and 0.00 for those runs of the unconstrained model.

Overall, these estimates imply strong and asymmetric effects of market concentration on adjustment speeds, in contrast to the negative but inconsequential dynamic effects estimated in the symmetric model. Banks in more concentrated markets, in comparison with the average bank, are *slower* to adjust deposit interest rates *upward* when below their estimated long-run equilibrium level; furthermore, at least in the case of MMDAs, they are also *faster* to adjust interest rates *downward* when they are above equilibrium. Thus, when market interest rates fluctuate in either direction, the adjustment behavior of banks in concentrated markets seems to allow them to extract more surplus from depositors than banks in less concentrated markets.

Estimates of the effect of concentration on the equilibrium markup, β_{HERF}, also differ from the symmetric model estimates. For CDs, the negative equilibrium effect in the asymmetric model is smaller in absolute value by about 20–40 percent. In the case of MMDAs the negative equilibrium effect shrinks by about 50–75 percent. Thus, the symmetry restriction not only obscures the effect of HERF on adjustment, but exaggerates the negative effect of HERF on long-run equilibrium. If adjustment is much slower from below in concentrated markets, then, on average, many more observations in those markets should be classified as below- versus above-equilibrium observations. The symmetry restriction on adjustment behavior creates a negative bias in the long-run slope coefficient by forcing a more even distribution of these observations above and below the equilibrium relationship.

A better idea of the meaning and magnitude of these estimates is conveyed by calculations given in Table V. Again, $\overline{\beta}$ is the estimated proportional markup for the bank with the mean characteristics. It is about 95 percent (of the T-Bill rate) in the case of CDs, and 90 percent in the case of MMDAs. The second row of the table gives the effect of a one-standard-deviation increase in the index of concentration on the equilibrium markup. For CDs and MMDAs, respectively, such a change lowers the equilibrium markup by about 1 percent and ½ percent of the total markup. In other

TABLE V
EFFECTS OF CONCENTRATION ON EQUILIBRIUM MARKUPS AND ADJUSTMENT SPEEDS*

	6MCD, long (1)	6MCD, long (1′)	6MCD, short (2)	6MCD, short (2′)	MMDA, long (3)	MMDA, long (3′)	MMDA, short (4)	MMDA, short (4′)
	Effect on equilibrium markups							
$\bar{\beta}$	0.947	0.960	0.947	0.962	0.897	0.927	0.890	0.932
$\beta_{\mathrm{HERF}}\sigma_{\mathrm{HERF}}$	−0.011	−0.010	−0.012	−0.009	−0.006	−0.005	−0.005	−0.008
	Asymmetry explained by concentration							
$\bar{\lambda}_A - \bar{\lambda}_B$	0.190	0.268	0.135	0.236	0.225	0.318	0.228	0.359
$\bar{\lambda}_{\mathrm{HERF}(A-B)}$	0.058	0.053	0.073	0.129	0.112	0.117	0.106	0.083

*$\bar{\beta}$, $\bar{\lambda}_A$, and $\bar{\lambda}_B$ are the equilibrium markup and adjustment speeds for a bank with the average characteristics. $\bar{\lambda}_{\mathrm{HERF}(A-B)} \equiv (\lambda_{\mathrm{HERF}(A)} - \lambda_{\mathrm{HERF}(B)}) \cdot \overline{\mathrm{HERF}}$, the amount of the asymmetry in the mean bank's adjustment speed accounted for by mean concentration level. $\overline{\mathrm{HERF}}$ and σ_{HERF} denote the mean and standard deviation of HERF given in Table I.

words, for a T-Bill rate of 750 basis points, such an increase in concentration lowers the equilibrium CD rate by about seven to eight basis points, and the equilibrium MMDA rate by about four to five basis points.

The estimated effects of concentration on the average bank's adjustment speeds appear more substantial. $\bar{\lambda}_A - \bar{\lambda}_B$ is equal to the difference between the mean bank's adjustment speeds from above and below equilibrium. In comparison, the final row of Table V gives the estimated contribution of HERF (for a bank with the mean HERF) toward "explaining" the asymmetry in these mean adjustment speeds, denoted by $\bar{\lambda}_{\mathrm{HERF}(A-B)}$, and computed as $(\lambda_{\mathrm{HERF}(A)} - \lambda_{\mathrm{HERF}(B)}) \cdot \overline{\mathrm{HERF}}$. That is, what is *not* "explained" by HERF is the implied difference in the adjustment speeds of a competitive bank (HERF $\approx$ 0). Generally speaking, HERF appears to explain between 25–50 percent of the total asymmetry.

The qualitative similarity of the effect of concentration on CD and MMDA pricing is notable given the differences in these two deposit accounts. Interest rates on MMDAs are on average more sluggish than those on CDs, especially in the upward direction. A price change instituted on CDs affects only marginal accounts—new CDs issued or old ones rolled over—and represents a contractual commitment. In contrast, for MMDAs, a change in price amounts to a repricing of all accounts, and confers no explicit contractual commitment on yields even one week into the future.

While the cross-sectional variation in deposit yields may, in principle, reflect unobserved quality or cost differentials, there is no reason to believe that these differentials are correlated with concentration. Furthermore, even if they are correlated, such differentials would be unlikely to change at a frequency comparable to the frequency of changes in yield differentials across markets; it is hard to see why compensation for persistent unobserved differentials should be realized through complicated differences in price adjustment behavior. Finally, given the asymmetry, the price sluggishness associated with market concentration clearly does not play the role of risk transfer often associated with implicit contracting.

What, then, could explain this asymmetry in price stickiness? While it is beyond the scope of this paper to provide a formal theoretical analysis, we offer a little speculation. It seems unlikely that menu costs could be the whole story. Since the six-month T-bill series is approximately I(1), the menu cost theory would seem to imply that the average nonzero deposit rate movement ought to be larger than the average T-Bill movement. In our sample, however, the average absolute T-Bill movement is 37 basis points, while the average nonzero 6MCD (MMDA) movement is 32 (26) basis points.

Plausibly, one might argue that banks in more concentrated markets are able to maintain low deposit rates longer because of tacit collusion, perhaps supported by the threat of future price wars (as in Rotemberg and Saloner [1986]. As the gap between market rates and deposit rates gets large enough, however, collusion can no longer support such low deposit interest rates, and so the banks partially adjust. On the other hand, when interest rates are falling, implicit collusion would support a prompt response by such banks, that is, the immediate lowering of their deposit interest rates.

V. Conclusions

In summary, market concentration is found to have a significant and substantial negative impact on average (over time) deposit interest rates offered by banks on MMDAs and six-month CDs. The results from estimating a symmetric model of dynamic deposit rate adjustment suggest that concentration affects only the level of deposit rates, but not the dynamics. However, an analysis of possible asymmetries in the dynamics reveals that, on average, banks are quicker to adjust their deposit rates downward when above "equilibrium" than they are to adjust upward when below equilibrium. Deposit interest rates of banks in more concentrated markets are found to exhibit stronger asymmetries in their dynamics.

These results therefore support the traditional view that downward price rigidity (the analogy to upward interest rate rigidity) exists—and is associated with market power. While the magnitude and duration of the effects found may appear small, it should be emphasized that consumer deposit markets are probably among the more competitive customer markets. Liquid asset markets are ordinarily thought to be a much closer approximation to the Walrasian market-clearing ideal than ordinary consumer goods markets. While consumers face some costs of switching, the product is relatively homogeneous, and price information relatively cheap. Producers, i.e., banks, operate on a relatively small average margin; furthermore, menu costs (of changing price) are not unusually large. Finally, the range and distribution of concentration, by the usual measures, is not out of line with that of standard consumer goods industries. For all these reasons, the finding of a dynamic impact of market power in this market suggests that such effects may be considerably more important in standard consumer goods markets.

UNIVERSITY OF PENNSYLVANIA
FEDERAL RESERVE BOARD

REFERENCES

Arak, Marcelle, Steven Englander, and Eric M. P. Tang, "Credit Cycles and the Pricing of the Prime Rate," *Federal Reserve Bank of New York Quarterly Review,* VIII (1983), 12–18.

Ausubel, Lawrence M., "The Failure of Competition in the Credit Card Market," Banking Research Center Working Paper No. 153, Kellogg Graduate School of Management, Northwestern University, 1989.

Berger, Allen N., and Timothy H. Hannan, "The Price-Concentration Relationship in Banking," *Review of Economics and Statistics,* LXXI (1989), 291–99.

Bok, Derek C., and John T. Dunlop, *Labor and the American Community* (New York: Simon and Schuster, 1970).

Carlton, Dennis W., "The Rigidity of Prices," *American Economic Review,* LXXVI (1986), 637–58.

——, "The Theory and the Facts of How Markets Clear: Is Industrial Organization Valuable for Understanding Macroeconomics?" in *Handbook of Industrial Organization,* Richard Schmalensee and Robert Willig, eds. (New York, NY: Elsevier Science Publishers B.V., 1989).

Diamond, D. W., "Financial Intermediation and Delegated Monitoring," *Review of Economic Studies,* LI (1984), 393–414.

Diebold, Francis X., and Steven A. Sharpe, "Post-Deregulation Deposit Rate Pricing: The Multivariate Dynamics," *Journal of Business and Economic Statistics,* VIII (1990), 281–93.

Domowitz, Ian, Glenn R. Hubbard, and Bruce C. Petersen, "Business Cycles and the Relationship Between Concentration and Price-Cost Margins," *Rand Journal of Economics,* XVII (1986a), 1–17.

Domowitz, Ian, Glenn R. Hubbard, and Bruce C. Petersen, "The Intertemporal Stability of the Concentration-Margins Relationship," *Journal of Industrial Economics,* XXXV (1986b), 13–34.

Flannery, Mark J., "Retail Bank Deposits as Quasi-Fixed Factors of Production," *American Economic Review,* LXXII (1982), 527–35.
Forbes, Shawn M., and Lucille S. Mayne, "A Friction Model of the Prime," *Journal of Banking and Finance,* XIII (1989), 127–36.
Gilbert, R. Alton, "Bank Market Structure and Competition: A Survey," *Journal of Money, Credit and Banking,* XVI (1984), 617–44.
Hall, R. L., and C. J. Hitch, "Price Theory and Business Behavior," *Oxford Economic Papers,* II (1939), 91–124.
Hannan, Timothy H., and Allen N. Berger, "The Rigidity of Prices: Evidence from the Banking Industry," American Economic Review, LXXXI (1991), 938–45.
Maddala, G. S., *Limited-Dependent and Qualitative Variables in Econometrics* (Cambridge, United Kingdom: Cambridge University Press, 1983).
Means, G. C., "Industrial Prices and Their Relative Inflexibility," Senate Document 13, 74th Congress, 1st Session (Washington, DC: U. S. Government Printing Office, 1935).
Moore, George, Richard Porter, and David Small, "Modeling the Disaggregated Demands for M2 and M1 in the 1980s: The U. S. Experience," presented at the Federal Reserve Board Conference on Monetary Aggregates and Financial Sector Behavior, 1988.
Okun, Arthur, *Prices and Quantities: A Macroeconomic Analysis* (Washington, DC: The Brookings Institution, 1981).
Rotemberg, Julio, "The New Keynesian Microfoundations," *National Bureau of Economic Research Macroeconomics Annual,* Stanley Fischer, ed. (Cambridge, MA: The MIT Press, 1987), pp. 67–104.
____, and Garth Saloner, "A Supergame-Theoretic Model of Business Cycles and Price Wars During Booms," *American Economic Review,* LXXVI (1986), 390–407.
____, and ____, "The Relative Rigidity of Monopoly Pricing," *American Economic Review,* LXXVII (1987), 917–26.
Roth, H. L., "Has Deregulation Ruined M1 as a Policy Guide?" *Economic Review,* Federal Reserve Bank of Kansas City, 1987.
Rudebusch, Glenn D., "Examining Alternative Econometric Specifications of the Disequilibrium Model: An Empirical Study with Labor Market Data," Economic Activity Section Working Paper No. 64, Federal Reserve Board, 1986.
____, *The Estimation of Macroeconomic Disequilibrium Models with Regime Classification Information,* Lecture Notes in Economics and Mathematical Systems (Berlin: Springer-Verlag, 1987).
Scherer, F. M., *Industrial Market Structure and Economic Performance* (Chicago: Rand McNally, 1980).
Simpson, T. D., "Implications for Monetary Policy of Changes in the Financial System," *Brookings Papers on Economic Activity,* I (1984), 249–72.
Stigler, George J., and James K. Kindahl, *The Behavior of Industrial Prices* (New York: National Bureau of Economic Research, 1970).