Monetary policy shocks and long-term interest rates

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and David Marshall

When monetary policymakers act, what happens to bond yields? There are good theoretical reasons why shorter-term bond yields should be affected by monetary policy. Open market operations of the Federal Reserve System have an immediate effect on the federal funds rate, which is the interest rate charged for overnight interbank loans. Since short-term borrowing (such as a one-month loan) acts as a reasonably close substitute for overnight borrowing, an increase in the federal funds rate should be accompanied by an increase in other short-term interest rates. However, it is less clear why monetary policy should have a significant effect on five-, ten-, and 15-year bond yields. It seems doubtful that five-year loans are close substitutes for overnight borrowing. Yet, casual observation suggests that monetary policy actions are associated with changes in long-term bond yields.

Consider the bond market debacle of 1994. Publications ranging from *Barron's* to the *Los Angeles Times* argue that 1994 was the worst year for the bond market since the 1920s. In figure 1, we display the one-year holding period returns for zero-coupon bonds of four years, six years, and ten years in maturity.[1] (The vertical lines toward the right-hand side of each panel indicate January 1994.) If we exclude the volatile period from 1979–82 (when the Federal Reserve experimented with direct targeting of monetary aggregates), the one-year cumulative losses in late 1994 were among the worst of the postwar period. This collapse of bond prices took its toll on well-known bond investors: Michael Steinhart's hedge fund sustained losses of 30.5 percent in 1994, George Soros's fell 4.6 percent, and Julian Robertson's fell 8.7 percent—all coming off very strong performances in 1993.

At the same time, 1994 was a period of concerted monetary tightening. After a period during which the federal funds rate was exceptionally low and stable, the Federal Open Market Committee (FOMC) raised the funds rate rapidly. As shown in figure 2, the 18 months from mid-1992 through the end of 1993 were characterized by a federal funds rate near 3 percent, with very little variability. This period more closely resembles the mid-1960s than the more volatile 1970s and 1980s. From February 1994 through February 1995, the FOMC doubled its target for the funds rate from 3 percent to 6 percent in seven increments. Figure 2 shows that this sort of monetary tightening is hardly unusual (even excluding the 1979–82 period, when the federal funds rate was not the monetary policy instrument). Nonetheless, the congruence of these two events (the rapid tightening of monetary policy and the precipitous rise in long-term

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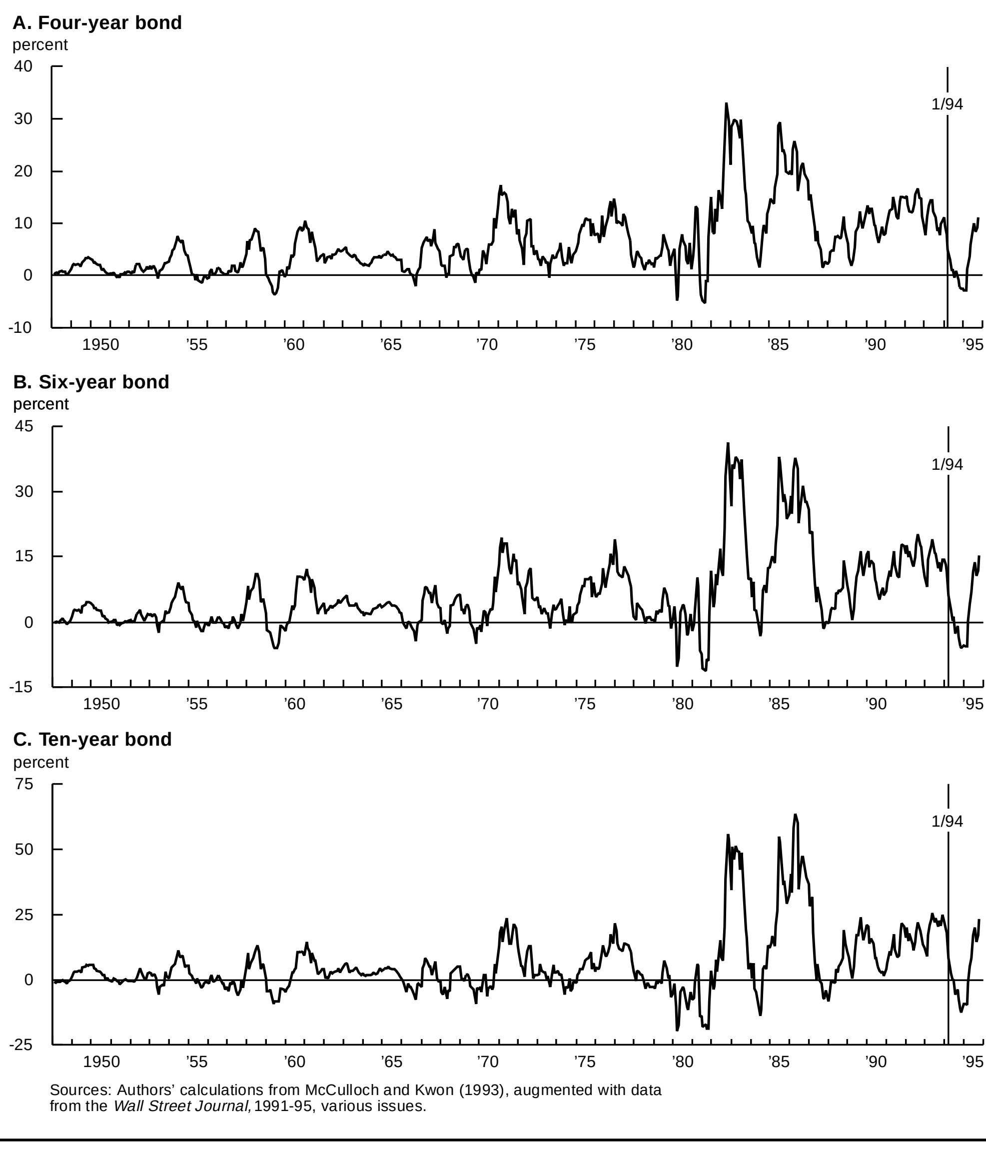
bond yields) led some to assert that the collapse in the bond market was policy induced. For example, the *Wall Street Journal* of December 13, 1995 graphically describes February 1994 as the month "when the Fed began raising short-term interest rates and set off the year's bond-market slaughter."

In this article, we will look at the relationship between monetary policy and long rates during the postwar period, and then apply what we learn to the extraordinary events of 1994. To examine how a monetary policy action (such as an increase in the federal funds rate) affects the yields of bonds with differing maturities, we must confront the problem of *causality*. For example, suppose we find that a tightening of monetary policy is associated with higher long-term bond yields. Can we then infer that tighter monetary policy *causes* higher yields? Not necessarily. It is generally believed that the FOMC tends to tighten monetary policy when there are indicators of future inflation. It is also believed that expectations of higher inflation tend to increase current bond yields. The positive correlation between tighter money and higher yields could

FIGURE 1

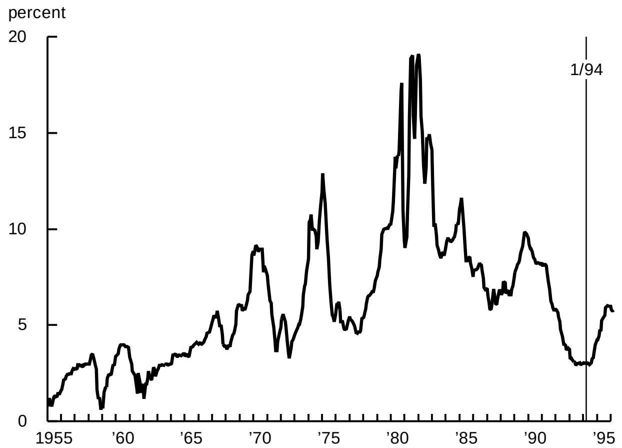
One-year holding period returns



Sources: Authors' calculations from McCulloch and Kwon (1993), augmented with data from the *Wall Street Journal,* 1991-95, various issues.

FIGURE 2

Federal funds rate



Source: Federal Reserve Board of Governors (1995).

We measure monetary policy by the level of the federal funds rate. We use the CEE framework to decompose changes in the funds rate into the feedback-rule component and the policy-shock component, and we ask how bond yields respond to an exogenous monetary policy shock. By focusing on the policy-shock component, we resolve the problem of ambiguous causality: Since the policy shock is exogenous by construction, causality can only flow *from* the policy shock *to* the bond yields (and to the other variables in the economy). However, this resolution is not without cost: We cannot ask how a change in the structure of the feedback rule itself would affect the behavior of long-term bond yields. The problem is that all observed economic relations are conditional on the particular feedback rule in place. [This is an application of the celebrated Lucas (1976) critique.]

be evidence that the Fed causes yields to increase when it tightens money, or it could be evidence that both the Fed action and the higher yields are jointly caused by forecasts of higher inflation.[2]

To help us disentangle the various possible directions of causality, we use a framework developed by Lawrence Christiano, Martin Eichenbaum, and Charles Evans in a series of working papers published by the Federal Reserve Bank of Chicago.[3] In the Christiano-Eichenbaum-Evans (CEE) framework, a clear distinction is made between the monetary authority's *feedback rule* and an *exogenous monetary policy shock.* The feedback rule relates policymakers' actions to the state of the economy.[4] In the example of the preceding paragraph, the "normal" reaction of the Fed to an increase in inflation expectations would be incorporated into the feedback rule. The exogenous monetary policy shock is defined as the deviation of actual policy from the feedback rule. We refer to these policy shocks as "exogenous" because, by construction, they do not respond in any systematic way to the economic variables that are included in the feedback rule. (If certain realizations of these variables systematically implied a higher-than-average or lower-than-average policy shock, then the rule is incompletely specified: Any systematic linkage between the policy-shock component and the feedback-rule component should have been loaded into the feedback rule in the first place.)

To explore the consequences of a change in the feedback rule, one would have to specify a model of the bond market at the level of investor preferences, monetary policy objectives, technological constraints, and market structure. We do not attempt this potentially useful but extremely difficult modeling task in this article.

Once we determine the response of bond yields to an exogenous monetary policy shock, we can look at the events of 1994 through this lens: (1) To what extent was the monetary tightening in 1994 an application of the FOMC's prevailing feedback rule, and to what extent did it reflect exogenous shocks to monetary policy?; and (2) To what extent did policy shocks affect long-term bond yields during this period? In particular, if there were *no* policy shocks (that is, if the monetary authority had followed the feedback rule *exactly*), would the increase in bond yields have been substantially reduced?

Below, we describe the CEE framework and how it is used to investigate the behavior of long-term bond yields. We then detail the average response of bond yields to exogenous monetary policy shocks. Our analysis indicates that these policy shocks have a substantial impact only on short-term bond yields; the impact on maturities longer than three years is quite small, and the impact on maturities longer than 15 years is insignificant. We consider two theoretical explanations for these results: the expectations hypothesis of the term structure and the Fisher hypothesis that movements in long-term bond yields reflect changes in expected inflation. We find that the response of long yields to exogenous monetary policy shocks closely follows the predictions of the expectations hypothesis, while the Fisher hypothesis explains very little. We then apply our methodology to the 1994 period.

A framework for analyzing the effects of monetary policy on bond yields

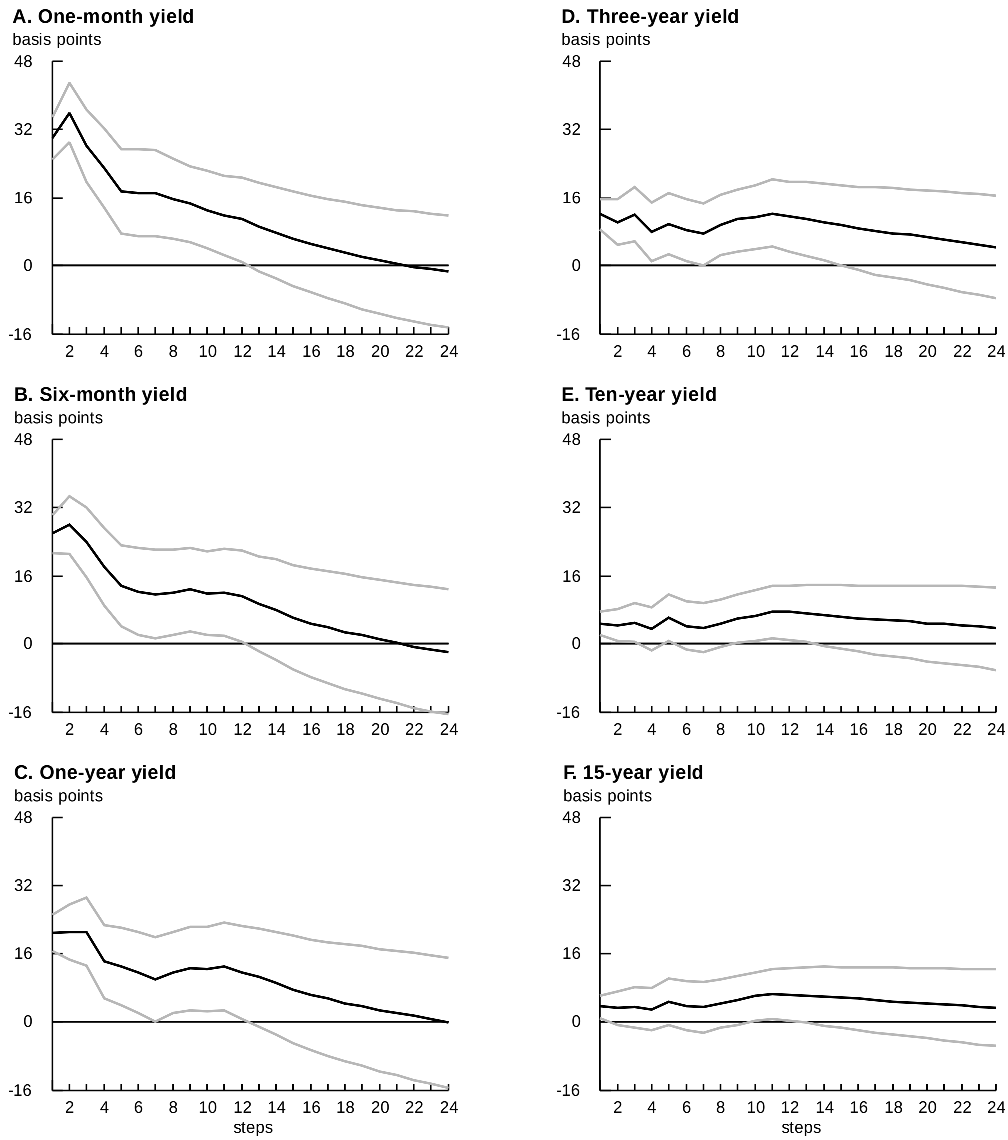
The model we use for exploring the effects of monetary policy shocks is the version of the CEE framework with monthly data discussed in Christiano, Eichenbaum, and Evans (1994b, section 5). We include four types of variables in our model. The first is the *monetary policy instrument*. We assume that this policy instrument is the federal funds rate. Christiano, Eichenbaum, and Evans (1994b) also explore the use of nonborrowed reserves as an alternative policy instrument. They obtain stronger results with the federal funds rate, but their results are fairly robust to the choice of instrument. The second type of variable is *contemporaneous inputs to the feedback rule*. We assume that this feedback rule incorporates contemporaneous values of the log of nonagricultural employment, as measured by the establishment survey (*EM*), the log of the price level, as measured by the personal consumption expenditure deflator (*PCED)*, and the change in an index of sensitive materials prices (*CHGSMP*).[5] We use *EM* as a monthly indicator of real economic activity. We measure the price level by *PCED*, rather than by the consumer price index (CPI), because the CPI is a fixed-weight deflator. Christiano, Eichenbaum, and Evans (1994b) discuss certain anomalous patterns that emerge when a fixed-weight deflator is used to gauge the price level.[6] These patterns are less of a problem when a variable-weight measure of consumer prices, such as *PCED*, is used. Finally, the *CHGSMP* series is a good predictor of future inflationary pressure. Some such predictor must be included if we are to construct a plausible representation of the Fed's feedback rule.

The third type of variable we include is *the yield on a zero-coupon bond* with T periods to maturity (Y^T); we rotate, one at a time, through maturities from one month to 29 years.[7] We use yields on zero-coupon bonds to avoid complications associated with coupon payments. The yields from 1947 to 1991 are monthly data taken from J. Huston McCulloch and Heon-Chul Kwon (1993).[8] For the period 1991 through 1995, we use yields on Treasury STRIPS quoted in the *Wall Street Journal* for the first business day of each month. Finally, we include *additional explanatory variables for long-term yields*. In this category of variables, we use the log of nonborrowed reserves (*NBR*) and the log of total reserves (*TR*). We use these variables as measures of the demand for credit in the economy. In particular, the amount of nonborrowed reserves that must be injected or withdrawn to achieve a given federal funds target is determined by the price elasticity of demand for reserves. By including total reserves as well as nonborrowed reserves, we measure the component of reserve demand that is accommodated through the discount window.[9]

The resulting model includes seven individual variables: *EM*, *PCED*, *CHGSMP*, *FF*, *NBR*, *TR*, and Y^T. We assume that the monetary policymakers' feedback rule is a linear function of (1) contemporaneous values of EM_t, $PCED_t$, and $CHGSMP_t$, and (2) lagged values of *all* variables in the economy. That is, the Federal Reserve sets policy based on current economic activity (as measured by EM_t) and price movements (as measured by $PCED_t$ and $CHGSMP_t$), as well as the entire history of the economy. The policy decision, in turn, has a contemporaneous effect on reserves and bond yields and affects the future realizations of all variables in the economy. Some argument could be made for including interest rates in the feedback rule, but there is statistical and economic justification for modeling the influence in the other direction. Cook and Hahn (1989) find that even on a daily basis there is little evidence of systematic movements in interest rates prior to an announcement of a change in the federal funds rate, while there are systematic movements after such an announcement.

FIGURE 3

The effect of a federal funds shock



Note: For each bond, the black line traces the response path over 24 months following the shock. The colored lines above and below the response give the 95 percent confidence bands, computed by Monte Carlo simulation using 1,000 independent draws.

Sources: Calculations from authors' statistical model, using the following data series: U.S. Bureau of Labor Statistics—employment survey measurements of nonagricultural employment (*EM*); U.S. Bureau of Economic Analysis—personal consumption expenditure deflator (*PCED*) and index of sensitive materials prices (*CHGSMP*); McColloch and Kwon (1993) augmented with data from *Wall Street Journal* 1991-95, various issues (Y^T); and Federal Reserve Board—federal funds rate (*FF*), nonborrowed reserves (*NBR*), and total reserves (*TR*).

We estimate this linear feedback rule as part of a vector autoregression (VAR) system. Formally, the system consists of seven equations. Each equation in the system takes one of the seven variables to be its dependent variable. For each equation, the independent variables are lagged values of all seven variables. The feedback rule consists of the fitted equation for *FF*, plus a linear combination of the residuals from the equations for *EM*, *PCED*, and *CHGSMP*. The exogenous monetary policy shock is that portion of the residual in the *FF* equation that is not correlated with this estimated feedback rule. The technical appendix to this article describes in detail how we set up and estimate this VAR, and how we use the VAR to infer the exogenous policy-shock component of *FF*.

In addition to the exogenous monetary policy shock, our model incorporates exogenous shocks to the other six variables. That is, there are a total of seven shock processes that act as the fundamental exogenous driving processes in the economy. These exogenous processes are transformations of the residuals from our seven VAR regressions. In particular, the exogenous shocks are serially uncorrelated, and are constructed to be mutually uncorrelated. (The technical appendix describes how we can isolate the effects of these seven exogenous processes.) Unexpected movement in any variable in the economy must be attributable to the effect of one or more of these exogenous processes. Below, we investigate how much of the unexpected movement in FF and Y^T can be attributed to the exogenous shocks to each of the seven variables in the model.

The response of bond yields to exogenous monetary policy shocks

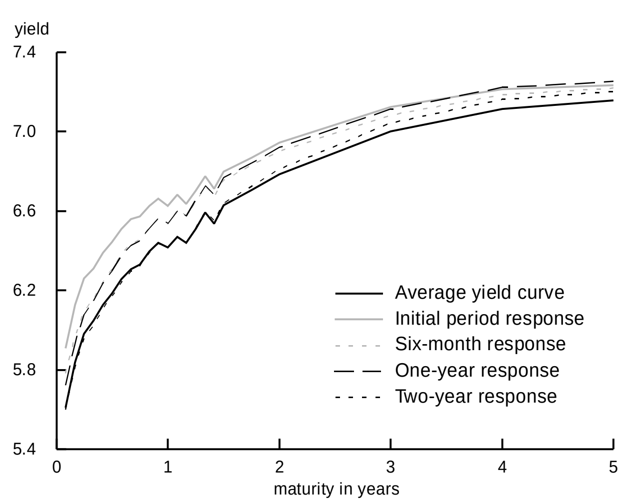
Figure 3 plots the estimated responses of bond yields to a one-standard-deviation exogenous monetary policy shock. This corresponds to an increase in the federal funds rate of approximately 50 basis points.[10] We display the responses for bond maturities of one month, six months, one year, three years, ten years, and 15 years. The colored lines delineate 95 percent confidence interval bands.[11] The plots trace the responses over 24 months.

A 50-basis-point federal funds shock increases the one-month rate by approximately 30 basis points in the period when the shock occurs. This response is highly significant statistically. The one-month rate continues to climb in the following period, and then falls, with the effect of the shock completely attenuated after 21 months. The six-month and one-year rates display qualitatively similar response patterns, although the magnitude of the response decreases for the longer-term bonds. When we move to longer-term bonds, the initial effect diminishes substantially as maturity increases: The initial response of the three-year bond is only 12 basis points, and the responses of the ten- and 15-year bonds are each less than 5 basis points. According to the point estimates, the response of the longer-term bonds appears more persistent than that of the shorter-term bonds. However, this apparent persistence is not statistically significant: The initial response for the ten- and 15-year bonds is barely significant; the response to a federal funds shock of all bonds longer than 15 years is insignificant at the 5 percent marginal significance level. For all maturities, the response is insignificant by one year. Interestingly, these results are roughly comparable to Cook and Hahn's (1989) estimates of the effects on interest rates of a publicly announced change in the federal funds rate. They find that in response to a 100-basis-point increase, short rates rise about 50 basis points, while long rates rise about 10 basis points.

The results are straightforward: There is a significant and relatively large effect on the short rates, with a decreasing, less significant effect at longer maturities. The effect on the term structure can perhaps be seen more easily by plotting the effect of a contractionary shock on the yield curve. The black line in figure 4 is the average yield curve from 1990 through 1995, for maturities up to five years. The remaining lines show our point estimates for the response of the yield curve to a one-standard-deviation exogenous monetary policy

FIGURE 4

The effect of a one-standard-deviation federal funds shock on the term structure



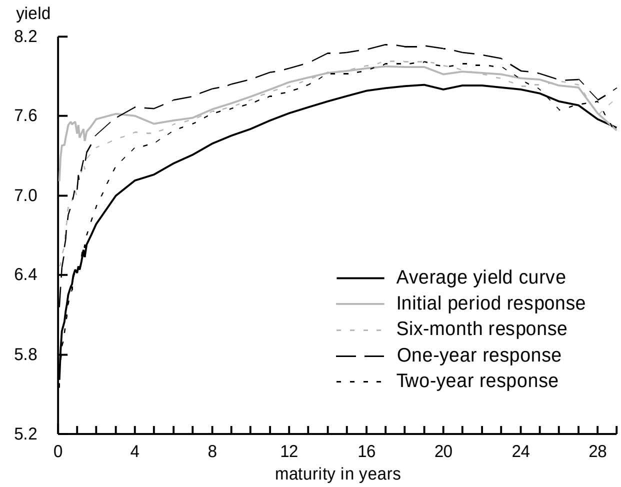
Sources: See figure 3.

shock after one month, six months, one year, and two years. To illustrate the qualitative patterns more clearly over a wider range of maturities, figure 5 displays a similar plot for a five-standard-deviation monetary shock, with maturities up to 29 years. These plots clearly show that the impact on the term structure is a rise in shorter rates, with the effect diminishing as maturities increase. In other words, a monetary policy shock raises the level, flattens the slope, and decreases the curvature of the yield curve.

Why do monetary policy shocks affect yields? What generates the observed response in yields of different maturities to a monetary policy shock? We consider two well-known hypotheses: the expectations hypothesis, which states that the long yield is an average of expected future short yields, and a version of the Fisher hypothesis, which states that changes in long yields are largely determined by changes in expected inflation.

FIGURE 5

The effect of a five-standard-deviation federal funds shock on term structure



Sources: See figure 3.

The expectations hypothesis

The expectations hypothesis can be written

$$1)\quad Y_t^T = \frac{1}{T}\sum_{i=0}^{t-1} E_t Y_{t+i}^1 + TP^T.$$

Equation 1 says that the long yield, Y_t^T, on a T-period zero-coupon bond is the average of expected future yields on one-period bonds over the next T periods, plus a time-invariant term premium, TP^T. The expectations hypothesis is attractive, because it implies that changes in forward interest rates should provide unbiased forecasts of changes in future spot rates. Unfortunately, tests of equation 1 using postwar U.S. data tend to decisively reject the hypothesis. For example, the equation implies that changes in the term spread $Y_t^T - Y_t^1$ should predict future yield changes $Y_{t+1}^{T-1} - Y_t^T$. That is, in the following regression

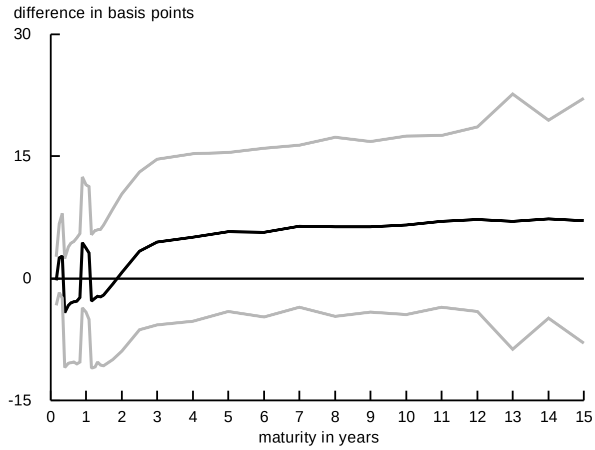
$$2)\quad Y_{t+1}^{T-1} - Y_t^T = a + b\,\frac{1}{T-1}\,[Y_t^T - Y_t^1] + e_{t+1}$$

the slope coefficient b should equal unity. Campbell and Shiller (1991) show that, for numerous data samples and numerous maturities T, this slope coefficient is significantly *negative*. McCallum (1994) suggests that the Campbell-Shiller regressions may be problematic econometrically in the presence of activist monetary policy. If the term premium TP^T displays only a small degree of time variation (so the expectations hypothesis holds approximately), but the monetary authority observes and responds to this time variation in TP^T, then e_{t+1} may be correlated with $Y_t^T - Y_t^1$. This could bias the slope coefficient b downward. McCallum gives examples where the resulting bias is sufficient to explain the Campbell-Shiller results.

In our impulse response functions, the expectations hypothesis would predict that the one-step-ahead response of Y_t^T should equal the average of the first T-period-ahead responses of the short rate Y_t^1. The Campbell-Shiller results suggest that the expectations hypothesis may perform poorly as an explanation of our impulse responses. On the other hand, our framework may not be vulnerable to McCallum's critique, since we model monetary policy explicitly. If the variables entering the feedback rule include those variables that shift the term premium,

FIGURE 6

Expectations hypothesis



Note: For maturities T ranging from two months through 15 years, the black line plots $resp_1(Y^T) - \frac{1}{T}\sum_{i=1}^{T} resp_{i'}(Y^1)$, where $resp_i(Y^j)$ denotes the response of Y^j_{t+i} (the yield on a j-period bond in period $t+i$) to a one-standard-deviation exogenous monetary policy shock in period t. The colored lines display 95 percent confidence bands.

Sources: See figure 3.

then our regressions will not display the McCallum bias.

In figure 6, we display the difference between the first-step response of Y_t^T and the average of the first T-step responses of Y_t^1, for T ranging from two months through 15 years. (The methodology used to construct the confidence intervals is described in the technical appendix.) According to this figure, the expectations hypothesis does a good job of explaining the impulse-response patterns. For all maturities, the difference between the first-period response of the long bond and the response predicted by the expectations hypothesis is less than 6 basis points, and is insignificant at the 5 percent marginal significance level.

The Fisher hypothesis

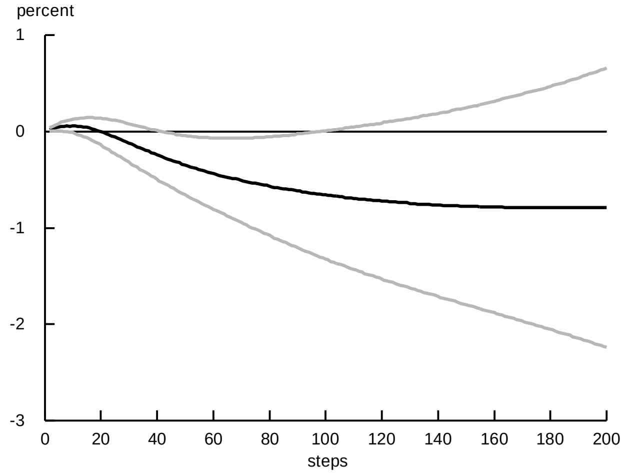
There is a school of thought that a good deal of the variation in very long-term bond yields is due to changes in expected inflation. An extreme version of this idea is the Fisher hypothesis, which asserts that the nominal bond yield Y_t^T should move, one for one, with changes in inflation expected over the life of the bond (that is, over the next T periods.)[12] Under this hypothesis, the *only* reason a monetary shock should affect long yields is because of its effect on the expected future price level: The first-period response of Y_t^T should equal the T-period-ahead response of the price level $PCED_t$.

There is substantial evidence against a literal one-to-one relationship between changes in expected inflation and changes in shorter-term interest rates.[13] However, it is not implausible that fluctuations in expected inflation are reflected, at least in part, in longer-term bond yields. To investigate this idea within our framework, we ask how much of the response of long yields to a monetary shock can be explained by the corresponding response in expected inflation. That portion of the response that cannot be tied to expected inflation would be attributable to liquidity effects, of the type described in Christiano and Eichenbaum (1992).

We assume that the impulse response of the price level is a good proxy, under rational expectations, for expected inflation following a shock in monetary policy. In figure 7, we display the response of our measure of the price level, *PCED*, to a one-standard-deviation contractionary shock to monetary policy. In figure 8, we display the difference between the first-step response of Y_t^T and the T-step-ahead response of $PCED_t$, divided by T in years, for maturities T ranging from two months through fifteen years. Unlike the expectations hypothesis, the Fisher hypothesis offers little explanation for our impulse responses. For all maturities, the difference between the first-period response of the bond yields and the response predicted by the Fisher hypothesis is significantly different from zero. Furthermore, the point estimates of these differences are fairly large, between 10 and 20 basis points. To see the source of this failure, compare figure 7 with figure 3. Figure 7 displays the response of the price level *PCED* to a one-standard-deviation monetary policy shock, along with the 95 percent confidence intervals. Initially, a monetary contraction is followed by a small (barely

FIGURE 7

The effect of a fed funds shock on the price level



Note: The black line represents the response to an exogenous monetary policy shock. The colored lines represent 95 percent confidence bands.

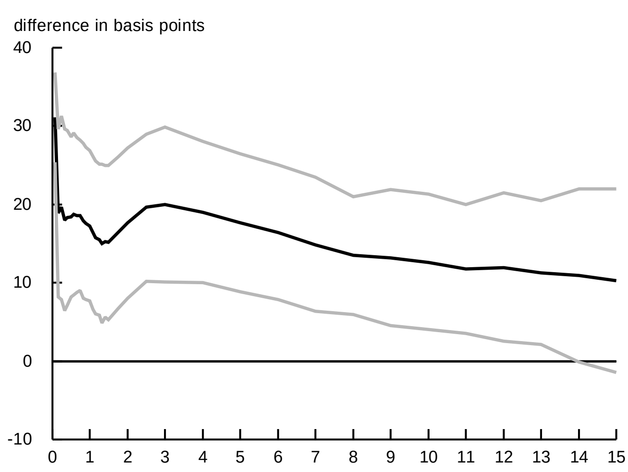
Sources: See figure 3.

significant) rise in the price level.[14] The price level eventually falls in response to a monetary policy shock. Under the Fisher hypothesis, this would imply a negative response of the longer-maturity yields to a monetary contraction. However, the estimated response of all yields to the monetary policy shock is positive. We find essentially no evidence that the response of long yields to an exogenous monetary shock is due to that shock's effect on expectations of future inflation.

In summary, we find that a contractionary exogenous shock to monetary policy has a strong upward impact on the one-month rate. One-month loans are a partial substitute for overnight borrowing, so it would be surprising if the one-month rate did not respond strongly to an increase in the federal funds rate. The impact of a shock to monetary policy on longer-bond yields declines with maturity, with this decline well explained by the expectations hypothesis. That is, the declining impact of a monetary policy shock on longer-maturity yields tracks the rate at which the response of the one-month yield attenuates. We find no evidence of an excessive response of long yields to monetary innovations. At the same time, changes in expected inflation do not appear to account for the observed responses.

FIGURE 8

The Fisher hypothesis



Note: For maturities T ranging from two months through 15 years, the black line plots $resp_1(Y^T) - \frac{1}{T(yrs)}\, resp_T\ (PECD)$, where for any variable X, $resp_i\ (X)$ denotes the response of X_{t+i} to a one-standard-deviation exogenous shock to monetary policy in period t. The colored lines display 95 percent confidence bands.

Sources: See figure 3.

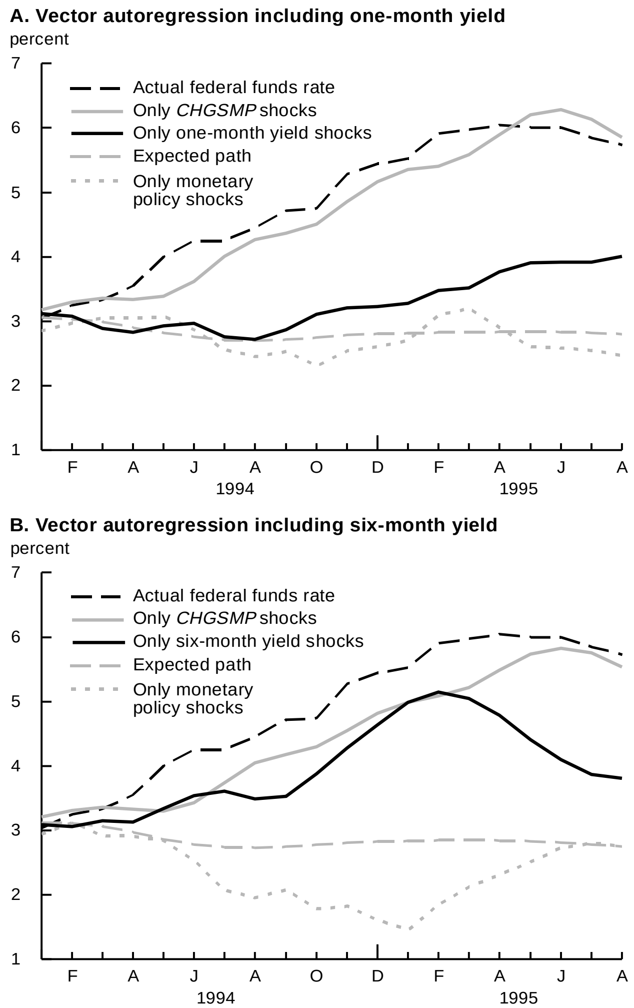
Monetary policy and bond yields in 1994

We use the results from the model to examine the behavior of monetary policy and the bond markets in 1994. Taking the VAR estimates as given, we decompose the movement of the federal funds rate and bond yields into the following: (1) the expected path, given information known in December 1993; (2) the unexpected movement attributable to the exogenous monetary policy shocks; and (3) the unexpected movement attributable to exogenous shocks to the other variables in the economy.

We first look at the determinants of the federal funds rate. Panel A of figure 9 shows

FIGURE 9

Historical decomposition of federal funds rate



Sources: See figure 3.

feedback rule. Figure 9 shows the relative importance of these two components. According to panel A of figure 9, the policy shocks account for virtually none of the unexpected run-up in the federal funds rate. Panel B of figure 9 indicates the exogenous policy shocks actually pull the federal funds rate *below* the expected path. It follows, therefore, that the increase in the federal funds rate must be due to the workings of the feedback rule. In particular, figure 9 indicates that most of the movement in the federal funds rate above the baseline forecast represents a response of the feedback rule to unexpected increases in sensitive materials prices. In both panels of figure 9, the line giving the path of the federal funds rate that would have obtained if all shocks other than shocks to *CHGSMP* were set equal to zero is very close to the path actually observed.

Recall that lagged values of the bond yield enter the feedback rule for monetary policy. Figure 9 documents the effect of shocks to the bond yield on the path of the federal funds rate. With the one-month bond (panel A of the figure), the exogenous shocks to the one-month yield have a rather small effect on the funds rate. When the six-month yield is used (panel B), the exogenous shocks in the bond yield do tend to push the funds rate above the expected path, but this effect is largely offset by the estimated exogenous monetary policy shocks.[16] The contributions from the inputs to the monetary policy rule other than the bond yield and *CHGSMP* are relatively small, so they are not plotted in figure 9.

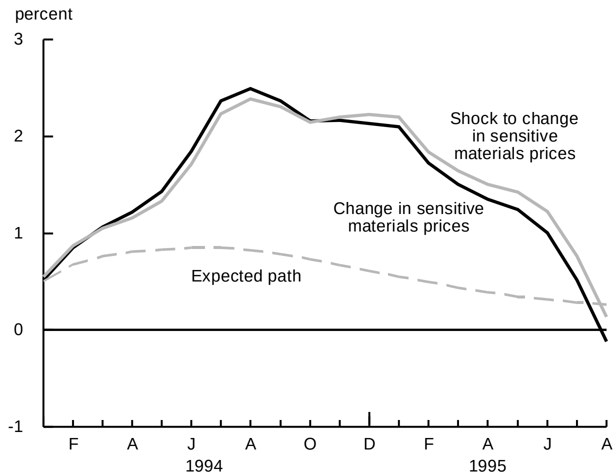
Our analysis indicates that the rise in the federal funds rate during 1994 and the first few months of 1995 was largely a mechanical response of the feedback rule to an increased

the decomposition of the federal funds rate when the VAR includes the one-month yield. Panel B shows the analogous decomposition when the VAR includes the six-month yield.[15] For both models, the expected path for the federal funds rate is flat. In contrast, the actual federal funds rate rises approximately 300 basis points from January 1994 through April 1995.

What accounts for this dramatic, unexpected tightening of monetary policy? By construction, the only sources of unexpected movements in the monetary policy are the exogenous policy shocks, and the effect of other economic shocks acting through the

FIGURE 10

Historical decomposition of sensitive materials prices



Sources: See figure 3.

threat of inflation. In our model, the monetary authority incorporates the series *CHGSMP* as a warning indicator of potential inflationary pressures. In 1994, this series took a pronounced and unexpected upswing. In figure 10, we display the actual *CHGSMP* series, along with the expected path of the series conditional on December 1993 information. Note that the growth rate in sensitive materials prices increases from 0.5 percent to 2.5 percent over the year, while the expected path does not even rise above 1 percent. Note that the line displaying the path the series would have taken if all shocks *except* the own-shocks to the *CHGSMP* series were set to zero closely tracks the actual series, implying that virtually all of this increase is attributable to the exogenous shocks to the *CHGSMP* series itself.

Were the increases in bond yields in 1994 and 1995 predictable? If not, why not? Figure 11 shows the historical decompositions for the one- and six-month yields, as well as the one-, three-, ten-, and 15-year yields. In all cases, the expected paths for the yields conditional on December 1993 information are flat. In contrast, all of these yields increased substantially during 1994. The increases range from approximately 300 basis points for the one-month yield to approximately 180 basis points for the 15-year yield. However, virtually none of this increase can be attributed to exogenous monetary policy shocks. In figure 11, we plot the path of each yield that would have obtained if the feedback rule were followed strictly. (That is, if the exogenous monetary shocks were all set equal to zero.) For each bond, the path is virtually unchanged.

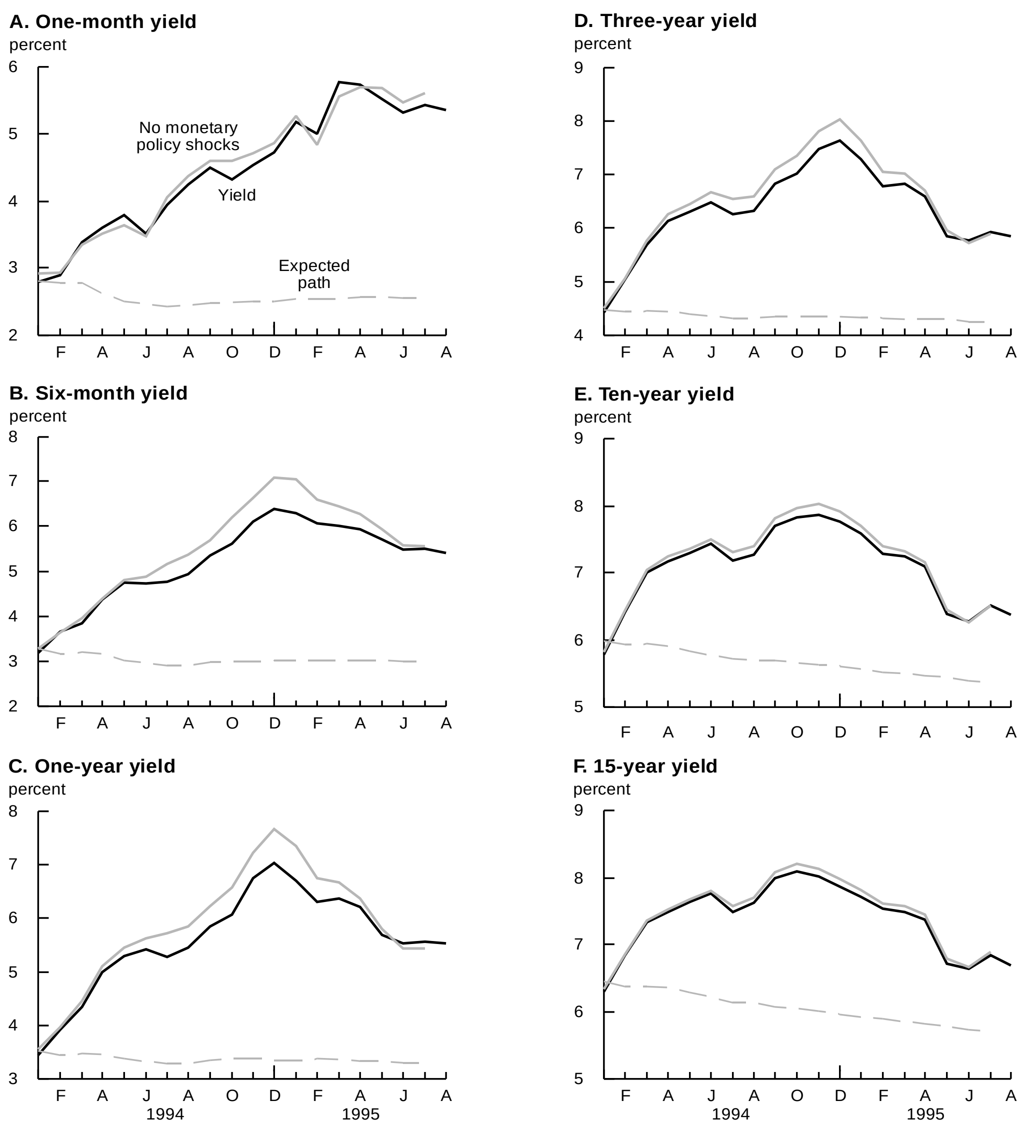
We can use our VAR model to explain the deviation of the bond yields from their expected paths. While some of these unexpected yield changes are a result of exogenous shocks to changes in sensitive materials prices (and, to a lesser degree, the remaining series in the model), for the most part, the unexpected movement in long bond yields is caused by exogenous shocks to the bond yields themselves. This is shown in figure 12. In each panel, the line tracing the path the bond yield would have taken if all shocks other than the own-shock to the yield itself were set to zero closely follows the movement in the bond yield. We interpret the exogenous shocks to the bond yields as shocks to financial markets that are unrelated to real economic activity (as measured by the employment variable EM_t), price changes, or monetary policy. The only other exogenous shock series that had a major impact on long-bond yields during this period is the shock to the change in sensitive materials prices. Our interpretation of the results in figure 12 is that the collapse in bond prices during 1994 was due, in part, to early warning signs of future inflation. However, this extraordinary movement in bond prices was largely due to factors that are unrelated to the economic or policy variables included in our model.

Conclusions

We find that there is a substantial response of one-month bond yields to an exogenous monetary policy shock, which dies out monotonically in about 20 months.

FIGURE 11

The effect of exogenous monetary policy shocks on bond yields, 1994–95



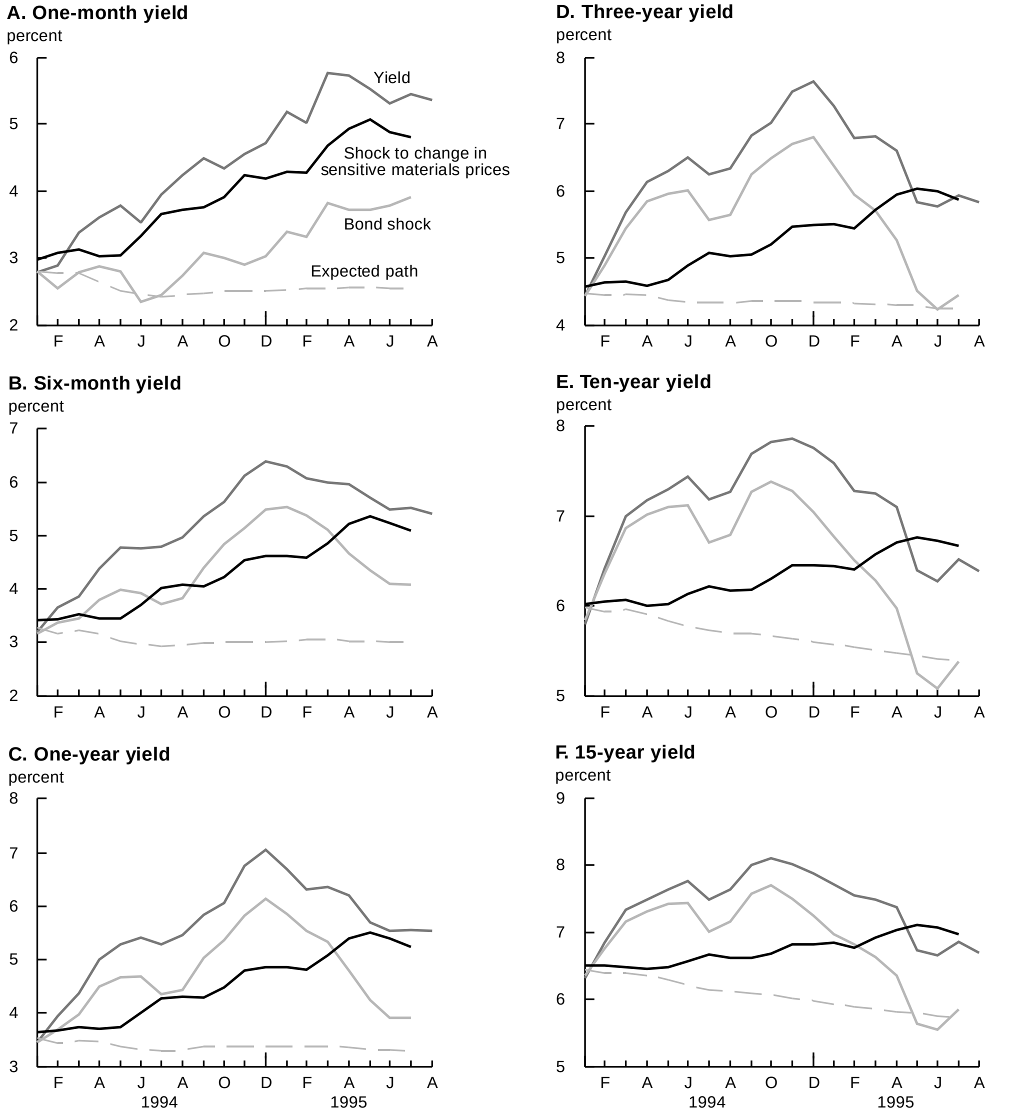
Sources: See figure 3.

Longer-term bond yields respond more or less as predicted by the expectations hypothesis: the initial month's response of a *T*-month bond's yield is approximately equal to the average of the first *T* months' response of the one-month bond. This pattern implies that longer-term bond yields have much weaker responses to an exogenous monetary shock. While these results are intuitive, they stand in sharp contrast to claims that long-bond yields react excessively to monetary policy innovations. We find no evidence that monetary policy shocks have any detectable effect on long-term bond yields.

When we apply these results to the dramatic events of 1994, we find no deviations from the general pattern. The substantial increase in long-term bond yields in 1994 cannot be attributed to exogenous monetary policy shocks. Indeed, the only evidence that might be interpreted as relating monetary policy to movements in long yields is the

FIGURE 12

The effect of exogenous shocks on bond yields, 1994–95



Sources: See figure 3.

impact of sensitive materials prices on both the federal funds rate and long yields. This could be evidence that increases in sensitive materials prices affected monetary policy *through the feedback rule*, and that this component of monetary policy might have had some impact on long yields. However, it is also possible that the change in sensitive materials prices affected long bond yields *directly*, rather than indirectly through the policy rule. For the reasons described in the introduction, there is no way we can disentangle these two pathways without a structural model.

TECHNICAL APPENDIX

To isolate exogenous monetary policy shocks, we use the vector autoregression (VAR) procedure developed by Christiano, Eichenbaum, and Evans (1994a, 1994b). Let Z_t denote the 7 x 1 vector of all variables in the model at date *t*. This vector includes the federal funds rate, which we assume is the monetary policy instrument, all inputs into the feedback rule, the long-bond yield being studied, and measures of nonborrowed reserves and total reserves. The order of the variables is:

A1) $Z_t = (EM_t, PCED_t, CHGSMP_t, FF_t, NBR_t, TR_t, Y_t^T)'$.

We assume that Z_t follows a sixth-order VAR:

A2) $Z_t = A_0 + A_1 Z_{t-1} + A_2 Z_{t-2} + \ldots + A_6 Z_{t-6} + u_t$,

where $A_i = 0,1, \ldots, 6$ are 7 x 7 coefficient matrices, and the 7 x 1 disturbance vector u_t is serially uncorrelated. We assume that the fundamental exogenous process that drives the economy is a 7 x 1 vector process $\{\varepsilon_t\}$ of serially uncorrelated shocks, with a covariance matrix equal to the identity matrix. The VAR disturbance vector u_t is a linear function of a vector ε_t of underlying economic shocks, as follows:

A3) $u_t = C\,\varepsilon_t$,

where the 7 x 7 matrix *C* is the unique lower-triangular decomposition of the covariance matrix of u_t:

A4) $CC' = E\,[\,u_t u_t'\,]$.

This structure implies that the *j*th element of u_t is correlated with the first *j* elements of ε_t, but is orthogonal to the remaining elements of ε_t.

In setting policy, the Federal Reserve both reacts to the economy and affects the economy; we use the VAR structure to capture these cross-directional relationships. We assume that the feedback rule can be written as a linear function Ψ defined over a vector Ω_t of variables observed at or before date *t*. That is, if we let FF_t denote the federal funds rate, then monetary policy is completely described by:

A5) $FF_t = \Psi(\Omega_t) + c_{4,4}\varepsilon_{4t}$,

where ε_{4t} is the fourth element of the fundamental shock vector ε_t, and $c_{4,4}$ is the (4,4)th element of the matrix *C*. (Recall that FF_t is the fourth element of Z_t.) In equation A5, $\Psi(\Omega_t)$ is the feedback-rule component of monetary policy, and $c_{4,4}\,\varepsilon_{4t}$ is the exogenous monetary policy shock. Since ε_{4t} has unit variance, $c_{4,4}$ is the standard deviation of this policy shock. Following Christiano, Eichenbaum, and Evans (1994), we model Ω_t as containing lagged values (dated *t* –1 and earlier) of *all* variables in the model, as well as time *t* values of those variables the monetary authority looks at contemporaneously in setting policy. In accordance with the assumptions of the feedback rule, an exogenous shock ε_{4t} to monetary policy cannot contemporaneously affect time *t* values of the elements of Ω_t. However, lagged values of ε_{4t} can affect the variables in Ω_t.

We incorporate equation A5 into the VAR structure A2 through A3. Variables *EM*, *PCED*, and *CHGSMP* are the contemporaneous inputs to the monetary feedback rule. These are the only components of Ω_t that are not determined prior to date *t*. The variables in the model that are *not* contemporaneous inputs to monetary policy but which do affect the long-yield under study are *NBR* and *TR*. Finally, the last element of Z_t is the long yield. With this structure, we can identify the right-hand side of equation A5 with the fourth equation in the VAR equation A2: $\Psi(\Omega_t)$ equals the fourth row of $A_0 + A_1 Z_{t-1} + A_2 Z_{t-2} + \ldots + A_6 Z_{t-6}$, plus $\Sigma_{i=1}^{3} c_{4i}\,\varepsilon_{it}$ (where c_{4i} denotes the $(4,i)$th element of matrix *C*, and ε_{it} denotes the *i*th element of ε_t). Note that FF_t is correlated with the first four elements of ε_t but is uncorrelated with the remaining elements of ε_t. By construction, the shock $c_{4,4}\varepsilon_{4t}$ to monetary policy is uncorrelated with Ω_t.

We estimate matrices A_i, $i = 0,1, \ldots, 6$ and *C* by ordinary least squares. The response of any variable in Z_t to an impulse in any element of the fundamental shock vector ε_t can then be computed by using equations A2 and A3.

The standard-error bands in figures 3, 7, and 8 are computed by taking 1,000 random draws from the asymptotic distribution of A_0, A_1, ... , A_6, C, and, for each draw, computing the statistic whose standard error is desired. The reported standard-error bands give the point estimate plus or minus 1.96 times the statistic's standard error across the 1,000 random draws.

To generate Monte Carlo standard-error bands in figure 6, our test of the expectations hypothesis, we must estimate an eight-variable VAR rather than the seven-variable VAR described in the text. The first six variables are unchanged; the last two variables are the one-month yield and the T-month yield, for T ranging from two months through 29 years. That is, the VARs now include *EM*, *PCED*, *CHGSMP*, *FF*, *NBR*, *TR*, Y^1, and Y^T, $T > 1$. Thus, 48 VARs were estimated, each with a different maturity's yield as the eighth variable. We use this modified VAR to calculate within a single model the difference between the first step response of Y_t^T and the average of the first T-period ahead responses of the one-month rate. The standard errors are then computed using 1,000 Monte Carlo draws, as described in the preceding paragraph. Note that the point estimate of the difference can also be estimated using the results from the seven variable VARs, which offers a good check of the eight-variable VAR method. The results are robust.

NOTES

[1]We use zero-coupon bonds to avoid the ambiguous impact of coupons on bond-price fluctuations. In particular, the effect of interest rates on bond prices (and therefore on holding period returns) depends both on the bond's maturity and on its coupon rate. Two ten-year bonds with different coupon rates will respond differently to a given interest rate shock. The behavior of one-year holding period returns for coupon bonds with durations of four, six, and ten years would be approximated by the plots in figure 1.

[2]A third direction of causality would be that an exogenous increase in yields induces a tight-money response by the Fed.

[3]See Christiano, Eichenbaum, and Evans (1994a, 1994b), and Eichenbaum and Evans (1992).

[4]Our use of the term "feedback rule" follows Christiano, Eichenbaum, and Evans (1994b). It should be clear, however, that the feedback rule is not a "law" and that there are no penalties for deviating from it. Rather, the feedback rule should be thought of as a set of quantitative relations that summarize the policymakers' normal response to economic developments.

[5]The variable *CHGSMP* is constructed by the Bureau of Economic Analysis. It measures the change in a composite index based on two sensitive materials price series, the producer price index of 28 sensitive crude and intermediate materials and the spot market price index of industrial raw materials.

[6]In particular, the price level displays a sustained *rise* following a monetary contraction.

[7]In this study, Y^T always denotes the *continuously compounded* yield to maturity. If y^T is the simple yield, then the continuously compounded yield is defined as $log\ (1 + y^T)$.

[8]McCulloch and Kwon (1993) provide yields on zero-coupon bonds for maturities through 40 years, but because of significant missing data, only rates through 29 years are used in our analysis. There are rates for monthly maturities from one to 18 months, then quarterly to two years, then semiannually to three years, and then annually to 29 years. All rates are annual percentage returns on a continuously compounded basis and are derived from a tax-adjusted cubic spline discount function, as described in McCulloch (1975). A more detailed explanation can be found in McCulloch and Kwon (1993).

[9]Other than the bond yields, all data are from the Federal Reserve's macroeconomic database. The series are monthly from 1959–95 and are seasonally adjusted where appropriate.

[10]The precise magnitude of a one-standard-deviation shock depends on the particular model, as follows: one-month rate, 50-basis-point shock; six-month rate, 49-basis-point shock; one-year rate, 48-basis-point shock; three-year rate, 49-basis-point shock; ten-year rate, 53-basis-point shock; and 15-year rate, 53-basis-point shock.

[11]Standard-error bands were calculated using the Monte Carlo procedure outlined in Christiano, Eichenbaum, and Evans (1994), with 1,000 Monte Carlo draws. The technical appendix describes this procedure in greater detail.

[12]To our knowledge, Irving Fisher never made the assertion implied by the hypothesis bearing his name. Fisher did note that if two risk-free interest rates are denominated in terms of different numeraires, they could differ only by the difference between the rates-of-change in the value of the numeraire goods. To derive the "Fisher hypothesis," one must combine Fisher's insight with the strong hypothesis that the real risk-free rate is constant, or at least uncorrelated with the inflation rate.

[13]See Marshall (1992) and the references therein.

[14]This initial rise in the price level is somewhat counterintuitive. One explanation is that the Fed uses information to forecast inflation that we have not included in our model. Since monetary policy affects the price level with some delay, the initial effect of a monetary tightening is to provide information that the Fed is forecasting future inflation. If these forecasts are accurate, on average, the initial response of the price level will be to rise. See Eichenbaum (1992) and Sims (1992) for further discussion of this issue.

[15]A decomposition of the monetary policy instrument when the one-month yield is included differs from the decomposition that includes the six-month yield because these are two distinct models of the monetary policy rule. We find that the decompositions with yields longer than six months have the same qualitative behavior as the decomposition using the six-month yield.

[16]A similar pattern obtains for all maturities longer than six months. For these longer maturities, the shocks to the yield tend to pull the federal funds rate below the expected path after March 1995. However, the exogenous monetary policy shocks also offset this effect.

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Management efficiency in minority- and women-owned banks

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Studies of the differences in operating performance of minority- and nonminority-owned commercial banks date back to the 1970s and early 1980s.[1] The focal point of much of this research was to investigate the long-term viability of minority-owned institutions. Some studies investigated declining lending trends among minority institutions (Boorman and Kwast 1974 and Meinster and Elyasiani 1988), while others concerned the possible adverse consequences of these trends on the economic development of the inner cities (for example, Kwast and Black 1983). As more attention is devoted to economic development prospects in our nation's core urban centers, the question of what role minority-owned banks (and other specially designated banks, including those owned by women) might play in the economic development of these communities naturally arises.[2]

Studies comparing the economic performance of minority- and nonminority-owned banks, for the most part, have revealed that the minority-owned banks have tended to be smaller, somewhat less profitable, and more expenditure prone than comparable groups of nonminority banks (Colby 1993). In addition, earlier studies reported that minority-owned banks tended to operate with lower ratios of equity capital to assets, to employ more conservative asset portfolio management policies, and to post higher loan losses than their nonminority peers (Brimmer 1971, Boorman and Kwast 1974, Bates and Bradford 1980, and Kwast 1981).

In contrast to these negative findings, a more recent study by Meinster and Elyasiani (1988) found that minority-owned banks had significantly improved their capital ratios and decreased their holdings of liquid assets, while expanding their use of purchased funds. The authors also reported that there were no significant differences in the pricing and asset-liability management decisions in the overall financial performance of minority-owned banks compared with a sample of nonminority-owned banks. However, Meinster and Elyasiani observed that banks owned by African Americans continued to reflect the financial performance characteristics associated with minority-owned bank performance in the 1960s and 1970s.

Caution must be exercised when comparing minority-owned with nonminority-owned banks on the basis of broadly defined markets or locational attributes. Studies by Clair (1988), Hunter (1978), and Mehdian and Elyasiani (1992) suggest that only when the two sets of banks are operating in identical or very similar market areas (in terms of economic and demographic characteristics) with similar customer bases is it safe to attribute differences in operating performance to differences in ownership and/or customer ethnicity.

Given the inherent difficulty in constructing samples of minority- and nonminority-owned banks which serve identical market

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areas, it is not surprising to find mixed conclusions in the literature assessing the long-term viability of minority-owned banks as engines of community economic development.[3]

In this article, we follow an approach similar in spirit to that used by Mehdian and Elyasiani (1992) in conducting an analysis of the operating performance of minority- and women-owned banks and comparable nonminority-owned banks from the perspective of production efficiency.[4] Instead of simply comparing the operating performance of a distinct sample of minority- and women-owned banks with a distinct sample of nonminority-owned banks, we compare the operating performance of our minority and nonminority sample banks relative to a set of so-called best-practice banks. This set of best-practice banks, which can include all types of banks regardless of ownership, represents those institutions which produce their financial products and services at the lowest cost using the most efficient mix of productive inputs or factors of production. Thus, unlike the older literature which infers managerial inefficiencies for minority-owned banks from simple comparisons of financial ratios, this article measures such managerial inefficiencies directly from the banks' cost (production) functions. We are thus able to determine which banks—various categories of minority- or women-owned and nonminority-owned—are more efficiently managed.[5]

Much of the literature examining the performance of minority banks is descriptive or based on regression analyses which lack well-developed theoretical underpinnings. In this article, we use production theory and modern econometric procedures to extract information on managerial efficiency in the production of financial services. Essentially, we estimate a firm-specific management efficiency measure for each bank in our sample using a standard bank cost function. As suggested by the earlier literature comparing the operating performance of minority- and nonminority-owned banks, differences in management efficiency among our sample banks could be due to a host of factors. Differences in managerial efficiency could result from differences in operating strategies, organizational structures, primary market areas, or customer bases. Below, we attempt to identify some of the determinants of observed managerial inefficiencies in our sample banks.

The empirical approach

In carrying out our empirical analysis, we use the methodology developed by Aigner et al. (1977) and Meeusen and Broeck (1977)—the stochastic cost frontier approach (described briefly below)—to calculate a measure of production efficiency (an inefficiency score) for each bank in our sample. These scores are used to gain further insight into the determinants of inefficiency.

Following Aigner et al. (1977) and Meeusen and Broeck (1977), a firm's cost function, that is, the relationship among the firm's total cost of producing various products or services, the products or services themselves, and the prices of the inputs used to produce these products or services may be written as

$$1)\quad TC_f = f\,(Y_i, P_k) + \varepsilon_f \quad f = 1, ..., n,$$

where TC_f represents the firm's total costs, Y_i represents the various products or services produced by the firm, P_k represents the prices of the inputs used by the firm in the production of the products or services, and ε represents a random disturbance term which allows the cost function to vary stochastically, that is, it captures the fact that there is uncertainty regarding the level of total costs that will be incurred for given levels of production. The uncertainty in the cost function can be further decomposed in the following manner:

$$2)\quad \varepsilon_f = V_f + U_f.$$

In equation 2, V represents random uncontrollable factors that affect total costs (such as weather, luck, labor strikes, or machine performance). These factors (and their impact on costs) are assumed to be independent of each another. They are identically distributed as normal variates and the value of the error term in the cost relationship is, on average, equal to zero.

The U term in equation 2 represents firm-specific cost deviations or errors which are due to factors that are under the control of the management of the firm. Such factors include the quantity of labor, capital, or other inputs hired or employed in the production of the firm's products and services and the amount chosen to be produced.[6]

The stochastic frontier cost function approach maintains that managerial or controllable inefficiencies only increase costs above frontier or best-practice levels, and that the random fluctuations or uncontrollable factors can either increase or decrease costs. Since uncontrollable factors are assumed to be symmetrically distributed, the frontier of the cost function, $f(Y_i, P_k) + e$, is clearly stochastic. In practical terms, the U component of the error term in the cost function given by equation 2, representing managerial inefficiency, causes the cost of production to be above the frontier or best-practice levels. Jondrow et al. (1982) estimated a firm's relative inefficiency using the ratio of the variability of the U and V terms in equation 2, which is measured by the ratio of the standard deviation $Q = s_u / s_v$, where s_u and s_v are the standard deviations of U and V. Small values of Q imply that the uncontrollable factors dominate the controllable inefficiencies.

In summary, the stochastic frontier approach incorporates a two-component error structure—one being a controllable factor and the other a random uncontrollable component. The controllable component consists of factors controllable by management.[7]

The cost function

To estimate the error term in the cost function given by equation 2 and to calculate each bank's efficiency index, we statistically fitted an empirical cost function of the following form:

$$3)\quad lnTC_f = \alpha_0 + \Sigma\alpha_i\, lnY_i + ½\, \Sigma\Sigma a_{ij}\, lnY_i\, lnY_j + \Sigma\,\beta_k\, lnP_k + ½\, \Sigma\Sigma\, \beta_{kh}\, lnP_k\, lnP_h + \Sigma\Sigma\,\gamma_{ik}\, lnY_i\, lnP_k + \varepsilon_f,$$

where TC_f represents total costs, Y_i represents the ith output, P_k represents the price of the kth input, ε_f is the disturbance term, and ln represents the natural logarithm. The cost function in equation 3 is a standard translog cost function. In fitting this cost function, standard homogeneity and symmetry restrictions were imposed.[8]

The sample data and variable definitions

The data for each sample bank examined were obtained from commercial bank "Reports of condition and income" filed with bank regulators. Average data for the four

TABLE 1

Frequency distribution of sample banks

A. Minority-owned commercial banks

	African American	Women	Hispanic American	Asian American	Native American	Total
Total	35	5	21	29	5	**95**
National charter	11	3	10	11	3	**38**
State charter	24	2	11	18	2	**57**
Bank holding company	17	5	8	10	1	**41**
De novo banks	3	0	0	7	0	**10**
Federal Reserve member	13	4	11	14	3	**45**

B. Nonminority-owned commercial banks

Total	National charter	State charter	Bank holding company	*De novo*	Federal Reserve member
127	66	61	59	6	82

Source: Federal Reserve Board of Governors, "Report of condition and income 1992," Washington, DC, magnetic tape, (April 1994).

quarters of 1992 was used. The sample was composed of all minority and women's banks and a comparable sample of nonminority-owned banks operating in 1992. The selection of comparable nonminority banks was based on size, location, market served, and start-up date. Initially, a nonminority-owned bank of similar size, established in the same year, with its headquarters in the same city as each sample minority or women's bank was identified. In cases where comparable banks could not be located, we expanded the search to encompass the metropolitan statistical area (MSA) of the minority- or women-owned sample bank. If we were unable to find a match in the same MSA, we selected an institution from a similar MSA market within the same state. This selection procedure resulted in a total of 127 banks being classified as comparable nonminority institutions. Panels A and B of table 1 provide data on the characteristics of the groups of banks.

Variable definitions

In the empirical cost function in equation 3, total costs (*TC*) were defined to include all labor and physical capital expenses, as well as the interest expense incurred by the bank, that is, the total costs of inputs used to produce the bank's various outputs. Four outputs were included in the cost function and were measured as the dollar value of (1) all money market assets, Y_m; (2) commercial and industrial loans, Y_c; (3) other loans, Y_l; and (4) other bank outputs, Y_o, which were proxied by annual noninterest income service charges, excluding gains and losses on foreign exchange transactions.

Labor, physical capital, and funds (including deposits) were treated as inputs used in the production of bank assets. With respect to input prices, the price of labor, P_1, was calculated by dividing total salaries and fringe benefits by the number of full-time equivalent employees (including bank officers). The price of physical capital, P_2, was defined to be equal to the ratio of total expenses for premises and fixed assets to total assets. The price of funds, P_3, was computed by taking the ratio of total interest expense (paid on deposits, federal funds purchased, securities sold under agreements to repurchase, demand notes issued to the U.S. Treasury, mortgage indebtedness,

TABLE 2

Mean values of key ratios

	Non-minority	All minority	African American	Women	Hispanic American	Asian American	Native American
Commercial loans	12.41	11.92	11.16**	17.00**	11.71	25.31***	10.32
Residential mortgage loans	18.17	13.57***	13.89***	7.88**	11.57***	10.91**	12.30**
Liquid assets	33.17	36.07	35.78	41.19**	41.48**	23.59***	45.68**
Delinquent assets	1.51	1.46	1.49	1.03*	1.05*	2.04**	1.15
Time deposits	40.19	43.48	42.75	33.52*	44.09	48.02**	48.49**
Retail deposits	13.12	14.49	13.93	7.12**	11.78	11.91	9.48
Interest expenses	2.97	3.08*	2.98	3.04	3.09*	3.10*	2.87
Noninterest operating expenses	4.01	4.33***	4.92***	4.17**	4.72**	4.97***	4.57**
Return on assets	.554	.485	.681**	.948**	.821**	–.309***	.568
Return on equity	5.91	5.78	7.41***	9.39**	9.61***	–.023***	5.53
Equity	9.03	8.86	7.83	7.87	7.48	11.15*	8.62

***, **, and * are significantly different from nonminority banks at the 1 percent, 5 percent, and 10 percent levels, respectively.

Note: All ratios except return on assets and return on equity are relative to total assets.

Source: Federal Reserve Board of Governors (1994).

subordinated debts and debentures, and other borrowed money) to the sum of total funds.

Empirical results

Table 2 provides some key balance-sheet and income expenditure ratios for the sample banks in our study. When minority- and women-owned banks were grouped in one category, called all minority, their asset portfolios and financing strategies were similar to those of nonminority banks, for the most part, except for a lower ratio of residential mortgage loans to total assets. In addition, the two groups' mean return on assets (ROA) and mean return on equity (ROE) were not significantly different. However, while African-American-owned banks had almost identical asset and financial statistics to those of nonminority banks, other minority- and women-owned banks were quite different from nonminority banks. Women-owned banks, for example, had higher ratios of commercial loans and liquid assets to total assets than nonminority-owned banks, but lower ratios of residential mortgage loans to total assets. They also posted lower ratios of time deposits and retail deposits to total assets than nonminority banks. On the other hand, Asian-American-owned banks had higher ratios of commercial loans and delinquent assets to total assets than nonminority-owned banks, as well as higher ratios of time deposits to total deposits. These banks also posted lower ratios of residential mortgage loans and liquid assets to total assets than nonminority-owned banks. In terms of profitability, the Asian-American-owned banks experienced negative returns over the sample period, while the other minority- and women-owned banks showed positive returns.

The descriptive statistics also show a significant difference in both the interest and noninterest operating expense categories between the groups of banks. The minority- and women-owned banks posted significantly higher ratios of noninterest operating expenses to total assets than did the nonminority banks. With respect to the ratio of interest expenses to total assets, all minority-owned banks again posted significantly higher ratios. However, among the minority- and women-owned banks, only the Hispanic-American and Asian-American banks had higher ratios of interest expenses to total assets.

Table 3 presents statistics for the variables used to estimate the cost function in equation 3. The input prices of minority- and women-owned banks exhibited a mixed pattern compared with those of the nonminority banks.

TABLE 3

Means for variables used in translog cost function

	Nonminority	As a percent of assets	Minority	As a percent of assets
Cost function inputs				
Price of labor	35.79 (16.83)	—	31.99* (7.12)	—
Price of capital	2.54 (1.88)	—	3.17** (3.21)	—
Price of funds	.036 (.017)	—	.032 (.007)	—
Cost function outputs *(mil.)*				
All money market assets	29.15 (48.95)	26.11 (16.10)	33.81 (70.05)	28.36 (15.70)
Commercial/industrial loans	12.93 (49.21)	12.41 (9.91)	14.45 (37.38)	11.97 (7.89)
Other loans	18.11 (79.37)	17.97 (30.04)	16.52 (65.43)	11.06** (28.12)
Other bank products and services	1.85 (7.61)	1.75 (7.50)	1.66 (4.10)	1.45* (1.01)
Total assets *(mil.)*	**102.3 (192.7)**	—	**120.8* (23.3)**	—
Total costs *(mil.)*	**7.54 (12.99)**	**7.87 (6.57)**	**8.43* (14.68)**	**7.80* (1.80)**

**, * Difference in means significant at the 1 percent and 5 percent levels, respectively.
Note: Standard deviations are in parentheses.
Source: Federal Reserve Board of Governors (1994).

While the price of funds at all of the sample banks was similar, the prices paid for capital inputs by minority- and women-owned banks were significantly higher, on average, than those paid by the nonminority banks. On the other hand, the prices paid for labor inputs were significantly lower at the minority- and women-owned banks. Despite this difference, total measured costs were significantly higher at the minority- and women-owned banks. In terms of asset allocation, the nonminority banks had a higher percentage of assets in commercial and industrial loans, other loans, and other bank products and services, but operated with a lower percentage of assets in the money market category than did the minority- and women-owned banks.

Management inefficiency

Higher capital input prices at minority- and women-owned institutions relative to the control group suggest inefficiency, particularly in light of the more liquid asset portfolios held by the minority- and women-owned banks.

Using the parameter values and standard errors of the residuals obtained from estimating a normalized version of the translog cost function in equation 3, inefficiency scores for the sample banks were calculated. The descriptive statistics displayed in table 4 suggest that both groups of banks produced products and services at a higher cost than necessary, that is, a perfectly efficient bank would have an inefficiency index of zero. The average inefficiency score of the minority- and women-owned banks was higher (31.4 percent) than the average inefficiency score of the nonminority-owned banks (24.8 percent) and the difference was statistically significant at the 5 percent level. Thus, on average, it appears that the minority- and women-owned banks were relatively inefficient institutions.

Asian-American-owned banks experienced the highest level of inefficiency (36.2 percent), followed by African-American (34.8 percent), Hispanic-American (33.1 percent), and Native-American banks (32.0 percent). Banks owned by women were more efficient than any of the other minority-owned banks but less efficient than the average nonminority bank. The results also indicated that the holding company structure was the most efficient structure for the minority- and women-owned banks. This could be the result of difficulties encountered by minority- and women-owned banks that are not affiliated with holding companies in adapting customer and

TABLE 4

Inefficiency score for sample banks

		Inefficiency score		
	Mean	Standard deviation	Minimum	Maximum
Nonminority banks	.248	.192	.056	.914
Minority banks	.314*	.105	.068	.966
African-American banks	.348*	.093	.032	.902
Women's banks	.267*	.168	.041	.925
Hispanic-American banks	.331*	.126	.035	.936
Asian-American banks	.362**	.110	.069	.955
Native-American banks	.320*	.098	.046	.928
National chartered banks	.318*	.108	.037	.944
State chartered banks	.320*	.112	.050	.958
Bank holding companies	.302**	.083	.074	.903
De novo banks	.347*	.148	.062	.941
Federal Reserve institutions	.332*	.130	.048	.921
Combined sample	2.710	.182	.035	.966

**, * Significantly different from nonminority sample banks at the 1 and 5 percent levels, respectively.
Source: Federal Reserve Board of Governors (1994).

service delivery systems in unique markets. It could also be due simply to a lack of managerial experience at these banks.

The relationship between firm inefficiency and bank characteristics was estimated using the following Tobit regression model:[9]

$$U_i = a_o + b_1 MINORITY + b_2 LIQUID\ ASSET + b_3 COMMERCIAL\ LOAN + b_4 RETAIL\ DEPOSIT + b_5\ ASSET + b_6 BHC + b_7 DE\ NOVO + b_8 NATIONAL + b_9\ 3\text{-}FIRM + b_{10} FEDMEMB + e_i,$$

where U_i = individual bank's inefficiency score,

MINORITY = minority- or women-owned indicator variable (1 for minority- and women-owned banks and 0 otherwise),

LIQUID ASSET = ratio of liquid assets to total assets,

COMMERCIAL LOAN = ratio of commercial loans to total assets,

RETAIL DEPOSIT = ratio of retail deposits to total deposits,

ASSET = total assets,

BHC = bank holding company dummy (1 if the financial institution is some form of bank holding company and 0 otherwise),

DE NOVO = *de novo* banks (1 for banks established within the last three years and 0 otherwise),

NATIONAL = national or state charter (1 for national chartered and 0 for state chartered banks),

3-FIRM = three firm deposit concentration ratio of respective metropolitan statistical market, and

FEDMEMB = Federal Reserve membership (1 for members and 0 otherwise).

In examining the determinants of inefficiency among the sample banks, we included variables related to portfolio composition (*COMMERCIAL LOAN*) and liquidity (*LIQUID ASSET*), financing or funding sources (*RETAIL DEPOSIT*), organizational characteristics [for example, whether the bank was a member of the Federal Reserve System (*FEDMEMB*) or organized as a holding company (*BHC*)], charter type (*NATIONAL*), market concentration (*3-FIRM*), and whether the sample bank was a *de novo* bank (*DE NOVO*).

While it is difficult to state *a priori* how each of these factors will influence bank inefficiency, it seems reasonable to expect *de novo* banks to be less efficient than other banks, and banks operating in concentrated markets to be less efficient than those operating in very competitive markets.

The regression results presented in table 5 show that the coefficient on the minority/women ownership dummy variable was positive and statistically significant. This implies that these banks were less efficient than their nonminority counterparts. Lending in the commercial and industrial loan category was also found to be associated with higher levels of inefficiency, while the bank holding company organizational structure was found to be associated with lower levels of inefficiency. As was expected, newly established banks tended to be less efficient than other banks and banks operating in less competitive markets tended to be less efficient than banks operating in more competitive, less concentrated markets.

Conclusion

Management efficiency has always been an important topic in banking research. Previous studies comparing the operating performance of minority- and women-owned banks with that of nonminority banks often reached mixed conclusions. This may have been due to the difficulty of identifying groups of minority and nonminority banks that are comparable along such dimensions as size and customer base. This article reported on the results of research which examined differences in the operating performance of minority- and women-owned banks from the viewpoint of production efficiency. Instead of simply comparing the operating performance of a distinct sample of minority- and women-owned banks with a distinct sample of nonminority-owned banks, we compared the operating performance of all of our sample banks relative to a set of best-practice banks. This set of best-practice banks, including all types of sampled banks regardless of ownership ethnicity or gender, represents those institutions that produced their financial products and services at the lowest cost using the most efficient mix of productive inputs or factors of production. Thus, unlike the older literature which suggests managerial inefficiencies for minority-owned banks from simple

TABLE 5		
Regression analysis		
Dependent variable: inefficiency score		
Independent variables	Tobit	
	Coefficient	Standard error
Intercept	.149	.024**
Minority	.058	.032*
Liquid asset	–.188	.112
Commercial loan	.060	.036*
Retail deposit	.132	.097
Asset	–4.7E-6	6.7E-6
BHC	–.073	.038**
De novo	.149	.061**
National	.223	.157
Fedmemb	.045	.036
3-firm	.092	.039***
Equation	Chi-Square = 142.06 * d.f. = 211	

***, **, and * are significant at the 1 percent, 5 percent, and 10 percent levels, respectively.

Source: Federal Reserve Board of Governors (1994).

comparisons of financial ratios, we measured such managerial inefficiencies directly from the banks' cost (production) functions.

We examined the performance of a sample of minority- and women-owned and nonminority-owned banks operating during 1992.

The results of our analysis indicated that, on average, while banks from both the minority- and women-owned and the nonminority categories were inefficient, the average minority- or women-owned bank was significantly more inefficient than the average nonminority bank. Among the sampled minority- and women-owned banks, the women-owned banks were the most efficient. Banks owned by Asian Americans were the least efficient among the minority-owned banks, followed by banks owned by African Americans and Hispanic Americans, respectively. *De novo* status was found to be a key factor accounting for higher levels of inefficiency. One explanation for this finding could be the lack of experience at *de novo* banks in serving new markets and customer bases. Another factor found to be important in determining the level of inefficiency among the sampled banks was the level of market concentration. The less competitive and more concentrated the bank's local market, the higher its level of inefficiency.

NOTES

[1]In this article, minority-owned banks include those owned by African Americans, Hispanic Americans, Native Americans, and Asian Americans. For a summary of history and trends in minority ownership of commercial banks see Price (1990).

[2]The recent controversy surrounding the acquisition of Indecorp, a leading Chicago minority-owned bank by ShoreBank Corporation, a nonminority-owned bank known internationally for its development efforts, is a case in point. See Wilke (1995).

[3]In this regard, Dahl (1995) offers a methodology which can potentially resolve this sample matching problem and, thus, contribute to our understanding of the observed differences in the operating performance of minority- and nonminority-owned commercial banks.

[4]Meinster and Elyasiani (1988) analyzed the 1984 year-end performance of a sample of 80 minority and 80 nonminority banks using a nonparametric efficiency technique—data envelopment analysis—based on linear programming principles. This technique assumes that all deviations from the best-practice cost frontier—including those due to random uncontrollable factors—are due to inefficient management. The stochastic frontier cost function approach used in this article does not assign deviations from the frontier caused by random uncontrollable factors to inefficient management.

[5]Research to date suggests that differences in managerial ability to control costs or maximize revenues account for as much as 20 percent of banking costs, while scale and scope inefficiencies account for only about 5 percent of costs. Thus, it is important to determine if there are significant managerial efficiency differences among banks owned by different ethnic and gender groups to draw more useful conclusions on long-term viability issues. See Berger et al. (1993).

[6]This inefficiency term is derived from a zero-mean normal, $N(0,\sigma_u^2)$, distribution truncated below zero. See Aigner et al. (1977) for a discussion and derivation of the cost function and error term structure given in equation 2.

[7]See Cebenoyan, Cooperman, and Register (1993) for a related estimation technique applied to thrift institutions.

[8]Symmetry requires that $\alpha_{ij} = \beta_{ji}$ and $\alpha_{hk} = \beta_{kh}$. The duality of the firm's cost and production function was preserved by imposing the following conditions: $\Sigma\beta_k = 1$, $\Sigma\beta_{hk} = 0$, and $\Sigma\gamma_{ik} = 0$.

[9]The Tobit regression model was used to eliminate the possibility of biased ordinary least square estimates where the dependent variable and error terms in the regression format are truncated normal variables (Amemiya 1973).

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