MONETARY POLICY AND THE CROSS-SECTION OF EXPECTED STOCK RETURNS

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Abstract

Ample evidence shows that size and book-to-market equity explain significant cross-sectional variation in stock returns, whereas beta explains little or none of the variation. Recent studies also demonstrate that proxies for monetary stringency increase the explained variation in stock returns. We reexamine a three-factor model that includes beta, size, and book-to-market equity, while allowing monetary conditions to influence the relations between these risk factors and average stock returns. We find that ex-ante proxies for monetary stringency significantly influence the relations between stock returns and all three risk factors. Additionally, all three variables are found to contribute significantly to explaining cross-sectional returns in a three-factor model that includes the monetary sector.

JEL Classifications: E44, E52, G12

I. Introduction

The roles of beta and other firm characteristics in explaining the cross-sectional variation in stock returns are well documented. Two variables found to be prominent are size, as measured by the market value of equity, and book-to-market equity (e.g., Fama and French 1992; Chan, Jegadeesh, and Lakonishok 1995). In this article we reexamine the significance of beta, size, and book-to-market equity, while allowing the monetary sector to influence the roles the factors play in the cross-section of returns. The analysis focuses on beta, size, and book-to-market equity because of their rich history in the literature. In selecting these variables, size and book-to-market equity appear to dominate most (if not all) other variables in the empirical explanation of cross-sectional returns, and beta has a rich theoretical foundation, but lacks empirical support.

We would like to thank Matt Billett, Werner De Bondt, William Higbee, an anonymous referee, and the editor for helpful comments. Participants at the 1998 Southern Finance Association meeting, the 1998 Financial Management Association meeting, the Texas Tech University Finance Workshop, and the 7th Global Finance Conference also provided helpful critiques for which we are grateful. This article received the “Outstanding Paper in Investments” award at the 1998 Southern Finance Association meeting.
We seek to determine whether the premiums attached to these factors vary in a predictable manner as the concerns of investors shift with changing monetary conditions. Like Merton’s (1973) and Breeden’s (1979) efforts, our approach proposes intertemporal links between a real macroeconomic variable and asset prices. We suggest that monetary conditions are linked with business conditions and market participants’ expectations about future market conditions. If changes in monetary conditions can be anticipated and are intrinsically linked to current and expected changes in consumption opportunities, investment opportunities, and the availability of money, then risk premiums may vary with perceived monetary policy changes. The answer to this issue can provide an important piece to the larger puzzle surrounding the search for economic state variables in asset pricing models.

Motivation for the study also comes from recent research showing that the single-factor specification is inappropriate (e.g., Fama and French 1992, 1995; Jagannathan and Wang 1996) and other recent studies linking the monetary sector to expected security returns (Jensen, Mercer, and Johnson 1996; Patelis 1997; Thorbecke 1997). Jensen, Mercer, and Johnson (1996), for example, find that changes in monetary stringency substantially increase the explained variation in the time series of aggregate excess stock returns. Ferson and Harvey (1999) find that common proxies for time variation in returns (e.g., the default spread) are significant in the cross-section of returns. Our article differs from these studies because we focus on cross-sectional returns and the influence of monetary stringency on the premiums attached to cross-sectional factors.¹

II. Sample and Variable Descriptions

Sample

We include all firms reporting return data in the New York Stock Exchange (NYSE)/American Stock Exchange (AMEX)/Nasdaq file from the Center for Research in Security Prices (CRSP) and financial data in the Compustat file from 1965 through 1997. The sample of firms in Compustat is limited in earlier years,

¹ An important difference between the Ferson and Harvey (1999) study and our study is that we allow the influence of the Fama-French (1993) factors to vary specifically with changes in our proposed state variable (i.e., monetary stringency), whereas they do not. Ferson and Harvey find that the inclusion of their conditioning variables in return regressions drives out the Fama-French factors. However, when we allow the influence of the Fama-French factors to vary with our monetary conditions proxies, the Fama-French factors remain significant predictors of cross-sectional returns. Given the importance of the Ferson and Harvey findings, we included a conditioning variable from their study (the default premium, based on the work of Chan, Karceski, and Lakonishok (1998) and Jensen, Mercer, and Johnson (1996)) in our estimations to examine whether the inclusion would affect the results. Although not presented here, the main findings and conclusions of our work are not affected by its inclusion.
which makes the 1965 start date appropriate. The sample includes only domestic firms with ordinary common shares outstanding. This sample is selected in light of the discussion in Chan, Jegadeesh, and Lakonishok (1995), who suggest that selection bias is not a severe problem for such firms.

**Variable Descriptions**

The empirical analysis examines the relation between cross-sectional portfolio returns and variables shown to be prominent in stock pricing models. Specifically, the pricing variables considered include market equity (ME), book-to-market equity (BE/ME), beta, and a proxy for monetary policy.

**Portfolio Returns.** We form portfolios in July of year \( t \) based on stock sorts that use accounting data for all fiscal year-ends in calendar year \( t-1 \). Post-ranking equally weighted portfolio returns are calculated monthly for July of year \( t \) through June of year \( t+1 \), and the portfolios are formed annually at the end of every June. Fama and French (1992) suggest that forming portfolios in this manner represents a conservative approach, ensuring that all accounting data are available before portfolio formation and the measurement of returns. To avoid problems with survivorship bias, a firm included in the sample at the beginning of a return year remains in the sample through the month it is removed from the data file.\(^2\) The sample returns start in July 1965 and end in June 1997, providing 384 monthly returns for each portfolio.

**Pre-Ranking Beta.** One of the sort variables on which we form portfolios is pre-ranking beta. Each stock’s pre-ranking beta is calculated as the sum of the estimated slope coefficients from regressions of monthly stock returns on the contemporaneous and previous month’s market returns. The inclusion of the previous month’s market return, also used by Fama and French (1992), helps alleviate the problem of nonsynchronous trading that plagues smaller stocks and results in the underestimation of beta for small firms. The model estimations use five years of monthly returns (ending with the return for June of year \( t \)) and the value-weighted CRSP index as the market proxy.\(^3\)

**Post-Ranking Beta.** Following Fama and French (1992), we calculate a post-ranking beta for each of the 125 portfolios (discussed later) using the full sample of post-ranking returns. The post-ranking portfolio betas are used in the models estimated in the following sections.

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\(^2\) The final monthly return for a firm that is removed from the file is measured as the return generated over the number of days during that month. For example, a firm deleted from the CRSP file after reporting ten daily returns in August would report a monthly return equal to the ten-day return. This procedure eliminates the bias created by omitting any returns of a firm removed from the file.

\(^3\) Following Fama and French (1992), we estimate betas using twenty-four to sixty months of return data (as available) in the five years before July of year \( t \).
Market Equity (ME). ME (size) is the second variable used to sort stocks into portfolios and is measured as the product of the number of shares outstanding and the price per share, both measured at the end of June in year \( t \).

Book-to-Market Equity (BE/ME). Book-to-market equity is the third variable used to sort stocks into portfolios. In BE/ME, the market value of equity is calculated by multiplying shares outstanding by the price per share, both at calendar year-end \( t-1 \). The book value of common equity is measured at fiscal year-end \( t-1 \). We adopt this procedure following Fama and French (1992), who note limitations to using market and book values measured at different times (which occurs for firms that do not have a December fiscal year-end). However, Fama and French note that their procedure is preferable to the alternative approach and that it does not bias their results. Shares outstanding and price per share are obtained from the CRSP file, and the book value of equity is obtained from Compustat.

Measures of Monetary Policy

We rely on three objective criteria to separate the Federal Reserve’s (Fed) monetary policy stance into two broad categories. Following Jensen, Mercer, and Johnson (1996), we first proxy the stringency of monetary policy using changes in the Fed discount rate to characterize policy periods as either expansive or restrictive. A discount rate decrease initiates a period of expansive monetary policy, and a discount rate increase initiates a period of restrictive monetary policy. The Fed is assumed to be operating under the same basic policy (e.g., expansive), regardless of the number of consecutive rate changes in the same direction (e.g., all decreases), until the discount rate is changed in the opposite direction. Therefore, a discount rate change in the opposite direction from the previous change initiates a new monetary environment, whereas consecutive changes in the same direction are considered a continuation of the then-current stance.

Three primary advantages of relying on changes in the discount rate as the monetary indicator are: (a) rate changes are perceived to be exogenous signals of Fed actions that are easily interpreted, (b) rate changes are widely reported and occur relatively infrequently, and (c) rate changes are regarded as signaling or confirming monetary developments and possibly real output developments (Laurent 1988).

The second proxy is based on the federal funds rate, which is advocated by several researchers as an effective way to gauge the Fed’s policy stance (e.g., Cook and Hahn 1989; Bernanke and Blinder 1992; Thorbecke 1997). We rely on changes in the Fed’s federal funds rate target because these changes are viewed as policy induced. The federal funds rate itself fluctuates considerably because of unanticipated reserve shortages over the reserve cycle and developments in sweep

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4 We wish to thank Kenneth Kuttner for providing data on the federal funds rate target.
account arrangements (Bennett and Hilton 1997). Thus, daily movements in the federal funds rate occur absent changes in Fed policy.

The third alternative proxy we use is the Boschen and Mills measure (see Boschen and Mills 1995; Thorbecke 1997; for discussions of this measure). The Boschen and Mills measure is fundamentally different from changes in the discount rate and changes in the federal funds rate target, as it is a narrative-based measure. Specifically, Boschen and Mills rely on an examination of Federal Reserve Open Market Committee records and similar documents to assess the Fed’s monetary policy stance.

Although our full sample using the discount rate provides 384 monthly observations, the federal funds target is available for only 220 months. These are limited because an explicit federal funds target is not continuously available (the Fed did not target the federal funds rate throughout the entire period). Because Boschen and Mills derive their measure only through the end of 1991, we have 319 monthly observations for this measure. To provide a consistent examination with the results presented based on changes in the discount rate, the two alternative measures are used to classify the monetary environment as either expansive or restrictive.

The model estimation results obtained from each of the three alternative measures are consistent. Because of the similarity in results and the need for conciseness, the model estimation results using the federal funds proxy and the Boschen and Mills proxy are not reported but are available on request. The finding of a high degree of consistency across three alternative monetary policy measures that are based on different factors and are developed independently supports the robustness of our results.

There is an important distinction between our use of a monetary policy indicator and studies examining beta’s influence in “up and down market” conditions (e.g., Wiggins 1992; Bhardwaj and Brooks 1993; Pettengill, Sundaram, and Mathur 1995). Notably, these studies define their subperiods in an ex-post manner, using information that would not have been available at the time. Furthermore, we broaden the focus by examining the role of size and book-to-market equity in the explanation of returns.

To examine the influence of monetary conditions within a regression analysis, we use discount rate changes to derive a directional monetary policy indicator ($D$). Table 1 provides summary information on the monetary policy dummy variable. Over the thirty-two-year (384 month) sample period, there are forty-one discount rate increases and forty-three rate decreases. However, there are only nineteen rate changes that are in the opposite direction of the prior change. Because the start of the sample period (July 1965) does not coincide with a rate change, there are twenty expansive or restrictive policy periods.

We take a conservative approach in our treatment of the nineteen months in which the monetary environment changed (i.e., the discount rate was changed in the opposite direction of the prior change). Specifically, at the start of any month,
TABLE 1. Series of Consecutive, Same-Direction Changes in the Federal Reserve Discount Rate.

<table>
<thead>
<tr>
<th>Series</th>
<th>Increasing (I) or Decreasing (D)</th>
<th>Month and Year of First Rate Change</th>
<th>Rate Changes in Series</th>
<th>Monthly Observations in Series</th>
</tr>
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<tbody>
<tr>
<td>1</td>
<td>I</td>
<td>07/65</td>
<td>1</td>
<td>22</td>
</tr>
<tr>
<td>2</td>
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<td>04/67</td>
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<td>11/67</td>
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<td>9</td>
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<td>08/68</td>
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<td>4</td>
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<td>12/68</td>
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<td>23</td>
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<td>D</td>
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<td>D</td>
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<td>7</td>
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<tr>
<td>20</td>
<td>D</td>
<td>01/96</td>
<td>1</td>
<td>17</td>
</tr>
</tbody>
</table>

Note: Series of consecutive increases (decreases) in the discount rate define periods of restrictive (expansive) monetary policy.

Over the sample period there were forty-one rate increases and forty-three rate decreases.

The number of monthly observations in the full sample equals 384, with 190 observations following rate decreases and 194 observations following rate increases. Months that include the first rate change in a series are taken as the final month in the prior series. The first and last months in the sample are July 1965 and June 1997.

an investor could not have known with certainty whether a rate change would occur in that month. Therefore, we classify each of these nineteen months as the final month of the previous monetary environment. This procedure avoids using ex-post information in establishing the return interval. The overall sample consists of 384 total months, with 194 months in restrictive policy periods and 190 months in expansive policy periods. The directional monetary policy dummy variable $D$ takes a value of 1 (0) in each of the 194 (190) months during restrictive (expansive) policy periods.

III. Empirical Analysis

Portfolio Formation

We form portfolios using a triple-sort procedure based on individual firms’ pre-ranking $\beta$, ME, and BE/ME. The triple-sort allows us to isolate return patterns
associated with an individual variable by controlling for return variation driven by both of the other measures. At the end of every June we rank all stocks based on their pre-ranking betas and form $\beta$-ranked quintiles. Stocks within each $\beta$-ranked quintile are then ranked and sorted into quintiles based on ME, providing 25 $\beta$:ME-ranked portfolios. Finally, we subdivide each of the 25 $\beta$:ME-ranked portfolios into quintiles based on BE/ME, creating 125 $\beta$:ME:BE/ME-ranked portfolios.\footnote{All estimations in this article were repeated using sixty-four $\beta$ : ME : BE/ME-ranked portfolios (i.e., a $4 \times 4 \times 4$ sort rather than a $5 \times 5 \times 5$). Furthermore, all estimations were repeated using all possible alternative sort orders for the three variables. The results are not materially different.} Equally weighted monthly portfolio returns are calculated for the following twelve months (July of year $t$ through June of year $t+1$), and portfolios are re-formed at the end of every June.

The process used to form portfolios serves several purposes. First, it helps alleviate the errors-in-variables problem that plagues betas on individual firms (see Blume 1971, 1975). Second, the triple sort creates dispersion in each of the portfolio characteristics that we wish to examine, while controlling for variation in both the second and the third characteristics. Third, the technique tends to orthogonalize the three independent variables and thus reduces the effect of multicollinearity in the regression analysis.

Regression Evidence

We estimate the models using the generalized least squares (GLS) approach of Parks (1967) in a pooled cross-sectional, time-series setting to control for time-series and cross-sectional correlation and heteroskedasticity in the model residuals. To avoid potential nonstationarity in the independent variables, we follow the detrending procedure of Chan, Hamao, and Lakonishok (1991). Size and book-to-market equity are adjusted to remove their general increase over time by dividing each variable by its respective cross-sectional average from the previous June. Hereafter, these two variables, when discussed in the regression results, are detrended.\footnote{We also estimated the models using the unscaled variables. The results are not materially different.}

We estimate the following models by pooling data on the 125 portfolios, where the dependent variable is the mean monthly portfolio return and the three independent variables are post-ranking portfolio $\beta$, portfolio $\ln$(ME), and portfolio $\ln$(BE/ME):

\begin{align*}
R_{pt} &= \alpha + \gamma_1(\beta_{pt}) + \gamma_2(\ln(ME)_{pt}) + \gamma_3(\ln(BE/ME)_{pt}) + \varepsilon_{pt}. \quad (1) \\
R_{pt} &= \alpha + \gamma_1(\beta_{pt}) + \lambda_1(D_t^* \beta_{pt}) + \gamma_2(\ln(ME)_{pt}) + \lambda_2(D_t^* \ln(ME)_{pt}) \\
&\quad + \gamma_3(\ln(BE/ME)_{pt}) + \lambda_3(D_t^* \ln(BE/ME)_{pt}) + \alpha' D_t + \varepsilon_{it}. \quad (2)
\end{align*}

Model (1) is estimated over the full period (348 months, $n = 48,000$), over expansive

<table>
<thead>
<tr>
<th></th>
<th>( \beta )</th>
<th>( \ln(\text{ME}) )</th>
<th>( \ln(\text{BE/ME}) )</th>
</tr>
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<tbody>
<tr>
<td>(1)</td>
<td>0.07</td>
<td></td>
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<tr>
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<td>(2)</td>
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<td></td>
<td></td>
<td>(−4.59)</td>
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<td>(3)</td>
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<td></td>
<td></td>
<td></td>
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</tr>
<tr>
<td>(4)</td>
<td>0.11</td>
<td>-0.06</td>
<td>0.35</td>
</tr>
<tr>
<td></td>
<td>(0.47)</td>
<td>(−2.84)</td>
<td>(7.72)</td>
</tr>
</tbody>
</table>

Note: Slope estimates and \( t \)-statistics are presented from a generalized least squares estimation using pooled cross-sectional, time-series data on 125 \( \beta \):ME:BE/ME-sorted portfolios over 384 months (i.e., \( n = 48,000 \) observations in each regression). The dependent variable is the equally weighted monthly portfolio return, and the independent variables are the post-ranking portfolio \( \beta \), the portfolio \( \ln(\text{ME}) \), and the portfolio \( \ln(\text{BE/ME}) \). Individual firms’ pre-ranking \( \beta \)s are used in the portfolio sorts. Post-ranking portfolio \( \beta \)s are used in the regressions and are calculated using the full sample (384 months) of portfolio returns and the value-weighted New York Stock Exchange (NYSE)/American Stock Exchange (AMEX)/Nasdaq file from the Center for Research in Security Prices (CRSP). ME for individual firms equals the number of outstanding shares times price per share, both measured at the end of June of year \( t \). BE/ME for individual firms is book common equity at fiscal year-end \( t−1 \) divided by ME calculated at calendar year-end \( t−1 \). The universe of stocks includes all domestic firms with return data in the NYSE/AMEX/Nasdaq CRSP file and financial data in the Compustat annual industrial and research files.

policy months (190 months, \( n = 23,750 \)), and over restrictive policy months (194 months, \( n = 24,250 \)). The inclusion in model (2) of the monetary policy dummy variable \( D \) interaction terms allows us to formally test whether significant differences exist in the slope coefficients across policy periods. The intercept and intercept dummy coefficient estimates provide no economic information and therefore are not reported in the table results.

Table 2 presents the GLS estimation results from model (1) over the full period. The inferences on the roles of the three variables in all of the estimations are similar to those provided by Fama and French (1992, Table III).\(^{7}\) That is, the slope estimate on \( \ln(\text{BE/ME}) \) is positive and significant and the slope estimate on \( \ln(\text{ME}) \) is negative and significant, whereas the slope estimate on \( \beta \) is not significant at traditional levels (even in isolation).

Table 3 presents the GLS estimation results from model (1) when the data are separated into expansive and restrictive monetary policy periods. The results show substantial differences in the coefficient estimates across the two monetary environments. The full model supports the hypothesis that the monetary environment

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\(^{7}\) We also estimated the models using the Fama and MacBeth (1973) methodology, with results similar to Fama and French (1992).

<table>
<thead>
<tr>
<th></th>
<th>Expansive Monetary Policy</th>
<th>Restrictive Monetary Policy</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \beta )</td>
<td>0.68 (3.67)</td>
<td>-0.65 (3.23)</td>
</tr>
<tr>
<td>\ln(ME)</td>
<td>0.64 (15.39)</td>
<td>0.33 (7.45)</td>
</tr>
<tr>
<td>\ln(BE/ME)</td>
<td>-0.13 (-6.47)</td>
<td>-0.03 (-1.12)</td>
</tr>
</tbody>
</table>

Note: Slope estimates and \( t \)-statistics are presented from a generalized least squares estimation using pooled cross-sectional, time-series data on 125 \( \beta \):ME:BE/ME-sorted portfolios over all 190 expansive monetary policy months (i.e., \( n = 23,750 \) observations in each regression), and separately over all 194 restrictive monetary policy months (i.e., \( n = 24,250 \) observations in each regression). The dependent variable is the equally weighted monthly portfolio return, and the independent variables are the post-ranking portfolio \( \beta \), the portfolio \( \ln(ME) \), and the portfolio \( \ln(BE/ME) \). Individual firms’ pre-ranking \( \beta \)s are used in the portfolio sorts. Post-ranking portfolio \( \beta \)s are used in the regressions and are calculated using the full sample (384 months) of portfolio returns and the value-weighted New York Stock Exchange (NYSE)/American Stock Exchange (AMEX)/Nasdaq file from the Center for Research in Security Prices (CRSP). ME for individual firms equals the number of outstanding shares times price per share, both measured at the end of June of year \( t \). BE/ME for individual firms is book common equity at fiscal year-end \( t - 1 \) divided by ME calculated at calendar year-end \( t - 1 \). The universe of stocks includes all domestic firms with return data in the NYSE/AMEX/Nasdaq CRSP file and financial data in the Compustat annual industrial and research files.

Influences the relations between the three pricing variables and cross-sectional returns. In periods of expansive policy, the coefficient estimates on all three variables have the expected signs and are significant at traditional levels. In contrast, in restrictive policy periods the coefficient estimates on two of the three variables differ from expectations. Specifically, during restrictive monetary periods, the coefficient estimate on beta is negative and significant and the coefficient estimate on \( \ln(ME) \) is not significant.

When beta is isolated, there is, on average, a large positive premium associated with it in expansive monetary policy periods and a negative premium for beta risk in restrictive policy periods. This result is consistent with the evidence found in prior studies (e.g., Bhardwaj and Brooks 1993) that beta and returns are positively (negatively) related in “up” (“down”) markets, using ex-post information to form the subperiods. Given the positive and negative subperiod estimates, it is perhaps not surprising to find that beta on average has no significant role in explaining returns over the entire sample.

When size is isolated, a significant small firm premium exists only in periods of expansive monetary policy. The coefficient estimate in restrictive periods

<table>
<thead>
<tr>
<th></th>
<th>$\beta$</th>
<th>$D\beta$</th>
<th>ln(ME)</th>
<th>$D\ln(ME)$</th>
<th>ln(BE/ME)</th>
<th>$D\ln(BE/ME)$</th>
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Note: Slope estimates and $t$-statistics are presented from a generalized least squares estimation using pooled cross-sectional, time-series data on 125 $\beta$:ME:BE/ME-sorted portfolios over 384 months (i.e., $n = 48,000$ observations in each regression). The dependent variable is the equally weighted monthly portfolio return, and the independent variables are the post-ranking portfolio $\beta$, the portfolio ln(ME), the portfolio ln(BE/ME), a monetary policy dummy variable that takes a value of 1 (0) in restrictive (expansive) monetary policy months, and interaction terms. Individual firms' pre-ranking $\beta$s are used in the portfolio sorts. Post-ranking portfolio $\beta$s are used in the regressions and are calculated using the full sample (384 months) of portfolio returns and the value-weighted New York Stock Exchange (NYSE)/American Stock Exchange (AMEX)/Nasdaq file from the Center for Research in Security Prices (CRSP). ME for individual firms equals the number of outstanding shares times price per share, both measured at the end of June of year $t$. BE/ME for individual firms is book common equity at fiscal year-end $t−1$ divided by ME calculated at calendar year-end $t−1$. The universe of stocks includes all domestic firms with return data in the NYSE/AMEX/Nasdaq CRSP file and financial data in the Compustat annual industrial and research files.

is substantially smaller and not significant. When book-to-market is isolated, there is a significant premium on high book-to-market stocks in both expansive and restrictive policy periods. The premium is considerably smaller and significant at a lower level, however, in restrictive periods.

To provide statistical evidence on the differences in the slope estimates, Table 4 presents the results from the GLS estimation of model (2). The full model shows that all three variables are significantly related to stock returns in expansive policy periods. Moreover, the significance of the interaction dummies shows that $\beta$, ln(ME), and ln(BE/ME) relate differently to returns across monetary periods. Specifically, the risk premium associated with each of the three variables is significantly smaller during restrictive policy periods. This finding suggests that the well-documented small firm and low book-to-market effects are related to the monetary environment.

The estimates also show that all three variables, in isolation, have significant relations of the expected sign with stock returns in expansive policy periods. The interaction dummy variables show, however, that the slope estimates for all three variables are significantly different in expansive and restrictive policy environments. Consistent with the results in Table 3, the evidence indicates that the
return premiums associated with each of these variables in isolation appear to be significantly linked to monetary stringency.

A reasonable argument is that real returns should be examined to determine whether the security return patterns and risk factor differences reported in Table 4 are driven by inflation that may be linked to monetary policy. Therefore, we reestimate the models with real returns as the dependent variable. In particular, the monthly percentage change in the consumer price index (CPI) is subtracted from the monthly portfolio returns to derive real returns. The results obtained using real returns are almost identical to the results in Table 4, suggesting that the differences in the coefficient estimates across monetary conditions are not driven by differences in inflation. Because the real return estimation results are so similar, we do not present them here. They are, however, available on request.

**Contribution of the Extended Model**

Although the test statistics reported in Table 4 indicate that the monetary policy interaction terms are statistically significant, they do not show the contribution that monetary policy has on the explanatory power of the model. This section provides evidence on the improvement in the explanatory power obtained by including monetary conditions. Specifically, the explanatory power of model (1) is compared with the explanatory power of model (2) to determine the benefit associated with including monetary conditions.

Because we use a GLS approach in a pooled cross-sectional, time series setting, the $R^2$ is not part of the statistical output from the procedure. In general, when using GLS, the traditional $R^2$ lacks a useful interpretation relative to the original data. Therefore, we employ an alternative measure that relates to the original data and has been advocated as a goodness-of-fit indicator under GLS (see Buse 1973). Because models (1) and (2) use stock returns as the dependent variable, we assess the relative merits of each model by examining the correlation between actual stock returns and the returns predicted by each of the models. The correlation of the actual returns and the predicted returns from model (1) is 5.7%, whereas the corresponding correlation from model (2) is a much higher 15%. This indicates that the inclusion of monetary conditions in the model substantially increases its accuracy in explaining stock returns.

**Survivorship Bias**

We provide evidence on the robustness of the results to the survivorship bias discussed in Kothari, Shanken, and Sloan (1995). In 1978, Compustat expanded its database from coverage of 2,700 NYSE/AMEX/Nasdaq firms to about 6,000 firms. The potential bias occurs because up to five years of data were added retroactively for many of these added firms. To avoid this transition period and its potential effects, we examine the 1978–97 period separately.

<table>
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<tr>
<th></th>
<th>$\beta$</th>
<th>$D*\beta$</th>
<th>ln(ME)</th>
<th>$D*\ln(ME)$</th>
<th>ln(BE/ME)</th>
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<tr>
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<td>(-4.79)</td>
<td>(5.39)</td>
<td>(11.40)</td>
<td>(-5.40)</td>
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<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
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<td>0.62</td>
<td>-0.35</td>
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<tr>
<td></td>
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<td>(-2.17)</td>
<td>(4.10)</td>
<td>(10.34)</td>
<td>(-3.52)</td>
</tr>
</tbody>
</table>

Note: Slope estimates and $t$-statistics are presented from a generalized least squares estimation using pooled cross-sectional, time-series data on 125 $\beta$:ME:BE/ME-sorted portfolios over 228 months (i.e., $n = 28,500$ observations in each regression). The dependent variable is the equally weighted monthly portfolio return, and the independent variables are the post-ranking portfolio $\beta$, the portfolio ln(ME), the portfolio ln(BE/ME), a monetary policy dummy variable that takes a value of 1 (0) in restrictive (expansive) monetary policy months, and interaction terms. Individual firms’ pre-ranking $\beta$s are used in the portfolio sorts. Post-ranking portfolio $\beta$s are used in the regressions and are calculated using the full sample (384 months) of portfolio returns and the value-weighted New York Stock Exchange (NYSE)/American Stock Exchange (AMEX)/Nasdaq file from the Center for Research in Security Prices (CRSP). ME for individual firms equals the number of outstanding shares times price per share, both measured at the end of June of year $t$. BE/ME for individual firms is book common equity at fiscal year-end $t-1$ divided by ME calculated at calendar year-end $t-1$. The universe of stocks includes all domestic firms with return data in the NYSE/AMEX/Nasdaq CRSP file and financial data in the Compustat annual industrial and research files.

The findings in Table 5 show that the estimation results from 1978 through 1997 are consistent with the results for the full period. Specifically, all three variables are significant and each has a significantly different coefficient estimate between expansive and restrictive monetary periods. This consistency suggests the results are both robust over time and robust to survivorship bias. Although the general results do not change, we note that this later period is characterized by smaller beta and size premiums, and a larger book-to-market premium.

IV. Discussion of Results

The cumulative evidence indicates that the monetary environment significantly influences the relations between risk factors and average stock returns. During expansive monetary policy periods, beta has a positive and significant relation to stock returns, both in isolation and in models that include size and book-to-market equity. For reasons that are not made clear by the evidence presented here, beta is negatively related to returns in restrictive policy periods.

A plausible explanation for the negative-beta result may be provided by prior studies such as Wiggins (1992), Bhardwaj and Brooks (1993), and Pettengill, Sundaram, and Mathur (1995). These studies show in an ex-post setting that beta
risk is penalized in down-market conditions. Thus, it is conceivable that restrictive monetary conditions capture this risk, leading to our finding of a negative beta coefficient estimate in periods when the Fed is tightening. In this article we capture this risk using ex-ante monetary policy information, and we show it is consistent across three alternative policy proxies. It appears that the variation in the beta risk premium associated with monetary policy causes beta to be unimportant in explaining cross-sectional returns when the model does not control for monetary stringency.

Market capitalization has the expected negative relation to stock returns during expansive policy periods, both in isolation and in models that include beta and book-to-market equity. During restrictive policy periods, however, the coefficient estimate on market capitalization is insignificant in both the simple and the multiple regressions. Monetary stringency also appears to influence the relations between returns and book-to-market equity, but less so than the relations with beta and market equity. Book-to-market equity is positively related to returns in both expansive and restrictive policy periods, both in isolation and in multivariate regressions. The findings indicate, however, that the relations between book-to-market equity and returns differ across expansive versus restrictive periods, with the relation significantly weaker in restrictive periods.

The relative distress factor of Chen (1991) and Chan and Chen (1991), and supporting empirical work in Fama and French (1993, 1995) and Jensen, Johnson, and Mercer (1997), provide a reasonable explanation for these findings. Their work supports the view that small firms and high book-to-market firms are more susceptible to distress and depressed earnings in a weakening economy. During these times, which would coincide with expansive monetary policy from an accommodating Fed, investors would require higher risk premiums on these characteristics. Conversely, investors would require substantially lower risk premiums when the economy is at its strongest and smaller, entrepreneurial or pure-play firms are seen as less risky. These risk factors thus become less important and are not heavily “priced” during times when the Fed would more likely be leaning against a strong economy through a restrictive policy. Additionally, Gertler and Gilchrist (1994) and Thorbecke (1997) argue that the size premium may vary with monetary policy developments because small firms have more restricted capital sources during periods of tighter money and are thus affected more by changes in monetary conditions.

Overall, the evidence that the risk premiums vary with monetary policy developments suggests that asset pricing models should consider the influence of the monetary sector. The evidence also supports the view of Harvey (1989) and Jagannathan and Wang (1996) that, because risk premiums are likely to vary, analyses that do not account for this variation may wrongly conclude that particular pricing variables are unimportant in explaining average returns. The insignificance of beta in recent studies may be partly explained by the evidence provided here showing its negative relation to returns in periods of restrictive policy and positive relation in periods of expansive policy. Prior studies do not allow for variation linked to the monetary sector.
V. Summary and Conclusions

We reexamine the relations between the cross-section of stock returns and beta, size, and book-to-market equity. The primary interest is whether the risk premiums associated with each of the three pricing variables vary systematically with changing monetary conditions, and whether the explained variation in average returns is improved when the influence of the monetary sector is included.

We focus on testing whether the monetary sector, as a possible state variable, influences the relation between stock returns and beta, size, and book-to-market equity. We find that the risk premium on beta varies significantly across expansive versus restrictive monetary environments. This finding supports Jagannathan and Wang’s (1996) contention that the risk premium attached to beta should vary over time. Our results also indicate that proxies for monetary stringency capture a significant portion of the time variation in the beta premium. In the same fashion, the monetary sector has a significant influence on the risk premiums associated with firm size and to a lesser extent book-to-market equity.

Fama and French (1993, 1996) argue that a three-factor model that includes the excess return on a broad market index, the return premium of small over big firms, and the premium of high over low book-to-market equity firms serves as a good model to explain the returns on portfolios that are formed to capture these asset pricing anomalies. Our article provides additional support for their three-factor model. However, we find that the monetary sector adds an important dimension by allowing us to capture variation in the risk premiums associated with these pricing variables.

Fama and French (1996) note that they are unable to identify the state variables that tie their results to Merton’s (1973) intertemporal CAPM or Ross’s (1976) arbitrage pricing theory. Our findings suggest that monetary conditions may proxy for a consumption-investment state variable that is of special hedging concern to investors and may provide an important link to these multiperiod models. Overall, our evidence supports recent research efforts showing that the single-factor specification of the CAPM, developed in a single-period setting, is inappropriate and must be modified for intertemporal considerations. We suggest monetary policy is one such important consideration and provide a simple ex-ante measure of monetary stringency that captures intertemporal variation.

References