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Investor risk premia and real macroeconomic fluctuations

Michael E. Fuerst *

Department of Finance, 514 Jenkins Building, School of Business Administration, University of Miami, Coral Gables, FL 33124, United States

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Abstract

The premium for risk should have a significant impact on real economic activity: basic finance teaches corporate managers to evaluate new investment opportunities based on a risk-adjusted cost of capital. Moreover, many asset pricing models show this premium should vary with economic conditions. Firms seeking to invest in the midst of a recession when the premium for risk is high, for example, must offer a higher rate of return on new investments. By inducing changes in investment, the risk premium should affect new durable orders, employment, and other measures of economic performance. We measure the risk premia implied by the Fama–French (Fama, E.F., French, K.R., 1993. Common risk factors in the returns on stocks and bonds. Journal of Financial Economics 25, 3–56) asset pricing model and test whether innovations in the premia are informative about future real economic activity. Our results indicate that these shocks, especially those to the premium related to small firms, induce responses in the real economy similar to those from monetary policy disturbances.
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* Tel.: +1 305 284 5289; fax: +1 305 284 4800.
E-mail address: mfuerst@miami.edu

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1. Introduction

Should the cost of capital have a significant impact on real economic activity? Basic finance teaches corporate managers to evaluate new investment opportunities based on the net present value of expected future cash flows. As the standard finance course cautions, this discount rate must be adjusted for the risk of the new project. If managers of risky firms are cognizant of their investors' required compensation for risk, an increase in the risk premium should raise the "hurdle rate" managers use to evaluate new investments. Consequently, an increase in the risk premium could quash many investments already planned and reduce the number of new feasible projects. On a macroeconomic level, the consequences of a rise in the hurdle rate should include a reduction in new durable orders, employment, industrial production, and consumption. We test these hypotheses and analyze how changes in the premium investors demand to bear risk affects subsequent real economic activity.

Our work is related to research on the stock market's ability to predict future macroeconomic conditions (Barro, 1990; Fama, 1981; Lee, 1992). If a stock's value represents the present value of its expected future dividends as implied by Lucas (1978), rational investors should bid up and down stock values based on their expectations of firms' profits. As firm profitability is highly correlated with future economic prosperity, a procyclical relationship between real activity and the stock market should be seen. In our work, we take the study of this interaction in a new direction by recognizing that stock market value may be decomposed into two components: One component is the expected future disbursement of firm profits, i.e., dividends. The second is the risk-adjusted interest rate that discounts these cash flows. The stock market is therefore a confounded measure of time-variant risk-adjusted discount rates and future firm cash flows. We extract the risk-premium component and examine its impact on real economic activity.[1]

The work of Lettau and Ludvigson (2002) and Lamont (2000) also focuses on investment decisions and their relationship to the risk premium. The analysis of Lettau and Ludvigson uses Q theory and a consumption-wealth ratio to proxy for the future risk premium and then analyzes the link between this proxy and future long-term investment. Lamont sheds light on a puzzle in post-war data: investment and stock returns are negatively contemporaneously correlated. He tests the conjecture that firms need to plan investments. Actual investing occurs with a lag following a change in the discount rate and induces the observed negative correlation. Using surveys of firm investment *plans*, Lamont demonstrates that planned investment rises as the stock market rises and risky discount rates fall, just as financial economic reasoning would imply.

In contrast, we measure the dynamic multifactor risk premia directly using the Fama and French (1993) asset pricing model and test whether these premia have implications for future real economic activity including new durable goods orders and housing starts. One advantage of our approach is that it frames the question within the asset pricing literature, the results of which drive corporate managers' financial decision-making. The Capital Asset-Pricing Model (CAPM) and, more recently, the intertemporal CAPM

[1] According to the present-value model, the price of a stock equals its expected future cash flows—its dividends—discounted at a constant or time-varying discount rate. When the stock return is expected to be constant and the dividends are expected to grow at a constant rate, this stock valuation formula reduces to the well-known Gordon Growth Model, $P = \mathrm{Div}_1/(r - g)$. See Campbell et al. (1997, p. 253).

(ICAPM) and the Arbitrage Pricing Theory (APT) form the standard framework managers use to calculate the risk-adjusted cost of capital. As standard practice dictates, the decision to invest is then made by either comparing the investment's internal rate of return against this hurdle rate or calculating the investment's net present value by discounting at this rate. We explicitly extract the risk premium component of such a capital budgeting analysis. A second advantage is our use of an ICAPM framework (Fama and French view their model as a version of the ICAPM) wherein valid state variables must act as predictors of changing consumption–investment opportunities. A priori, such a model is well suited to the study of linkages between asset pricing and future macroeconomic fluctuations.[2]

In empirical tests of the CAPM, firm market capitalization or size is found to explain a significant portion of the required return (Banz, 1981). Hardouvelis and Wizman (1992) compare the variation in this required return by firm size over the business cycle and find a flight to quality: The cost of capital for small firms shows greater cyclicality than that for large firms. In asset-pricing terms, the risk premium for size shows strong countercyclical variation. Hardouvelis and Wizman conclude that this size effect may be a significant propagation mechanism of business cycles.

We extend this work and examine the time series behavior of the risk premia on the three factors of the Fama and French (1993, 1996) model. Similar to the model used by Hardouvelis and Wizman (1992), the Fama–French model includes a small firm state variable (SMB) in addition to the market factor of the CAPM. (Fama and French (1996) note that small firms load heavily on SMB.) We also draw parallels between this small firm risk premium and mechanisms of monetary policy transmission through both the balance sheets of borrowers and the credit channel (Bernanke, 1993; Gertler and Gilchrist, 1994). In both mechanisms, small firm access to capital plays a prominent role.

Moreover, the Fama–French model includes a state variable related to relative distress (HML) as measured by a (higher) book-to-market-equity ratio. Cochrane (2001) interprets relative distress in light of the findings of Heaton and Lucas (2000): The typical stockholder is the owner of a privately held business. Such businesses are particularly vulnerable to financial distress. To offset this risk, the average investor holds in his portfolio firms that are less vulnerable to distress. Hence, HML may capture an economic risk factor related to investors' *income* from *productive* activities.

Recent empirical research has directly linked the Fama–French factors for relative distress and size to future macroeconomic states. Liew and Vassalou (2000), Vassalou (2003) find that SMB and HML capture a risk factor related to future growth in GDP. In international markets, Liew and Vassalou show that the SMB and HML portfolios—and therefore stocks that weight heavily on these factors—do well in good times and poorly in bad times. This is the exact opposite of what the consumption risk-averse investor desires and justifies risk premia on these factors. Examination of the premia on these factors separately allows us to explore how risks related to the financial positions of small and vulnerable firms asymmetrically impact the macroeconomy.

The paper is organized as follows: Section 2 discusses the literature on the dynamics of the equity risk premium. Section 3 describes the econometric model and methodology. Section 4 relates our measures of the risk premia to the theoretical asset pricing models

[2] As Cochrane (2001) describes, drawing a distinction between an ICAPM and APT model in empirical work is usually difficult. We will often refer to factors (an APT term) and state variables (an ICAPM term) interchangeably.

described in Section 2. Section 5 describes the impulse-response function analysis. Section 6 relates our results to monetary policy mechanisms, and Section 7 concludes.

2. Dynamics of the risk premium

Several general equilibrium asset-pricing models endogenize the equity premium and allow it to vary as a function of past, current, and expected future consumption levels and volatility. For reasonable model parameterizations, the risk premium in Black (1990), Abel (1988) rises as current consumption falls, a consequence of a rising marginal utility of current consumption. The equilibrium factor model of Connor and Korajczyk (1989) also exhibits a risk premium that varies inversely with current output (under the assumption of constant absolute risk aversion and serially uncorrelated production). In addition, the models of Abel (1988, 1999) find a risk premium that rises as expected *future* consumption rises[3] and expected future consumption *volatility* increases. Subject to parameter restrictions, all these models imply a risk premium that varies countercyclically with the business cycle and even leads it. At business cycle peaks, when current consumption and output are high, investors have a relatively lower marginal utility for consumption and therefore require less compensation for bearing risk. At business cycle troughs, investors suffer from lower current consumption and anticipate higher future consumption levels and volatility (Kandel and Stambaugh, 1990). Therefore, the risk premium is at its greatest at the depths of a recession.

Consideration of habit formation[4] has allowed asset pricing models to better explain the unrealistic implications of earlier models including the well-documented equity premium puzzle of Mehra and Prescott (1985). In the Constantinides (1990), Campbell and Cochrane (1999) models, for example, current consumption enters the utility function net of a benchmark or habit level of consumption. Empirically, this implies that the equity risk premium is also a function of current and lagged consumption. In simulations of their model, Campbell and Cochrane find a slowly time-varying countercyclical risk premium.

Several empirical findings support the countercyclical and leading behavior of the risk premium (or, equivalently, procyclical and leading behavior of the stock market). In a very early study, Mennis (1955) finds that the Dow Jones Industrial Average led 18 out of 22 peaks and troughs in the business cycle over the period 1899–1949 and the Standard and Poor's Composite Index led 28 out of 34 turns in the business cycle from 1871 to 1949. In fact, real stock returns are positively related to and actually lead several real variables such

[3] The implication of these models that a rise in expected future consumption should lead to an increase in the current premium for risk may seem counterintuitive. Our intuition tells us that a rising risk premium quashes new capital investment and lowers future output and hence consumption. This apparent paradox can be explained as follows: If investors expect future consumption to increase, their desire to save for the future decreases, and consequently, the premium necessary to induce them to invest rises. Nevertheless, positive shocks to the risk premium increase the cost of capital, lower new investment, and lead to relatively lower realized future consumption. Therefore, expected future consumption and the risk premium should show positive contemporaneous correlation. However, realized future consumption and positive shocks to the current risk premium should show negative correlation. A parallel to monetary policy is insightful: if the monetary authorities expect a rise in the future rate of inflation, they may raise the fed funds rate. Nevertheless, innovations in the fed funds rate are contractionary and should lead to a reduction in realized inflation.

[4] Habit formation is the psychological phenomenon wherein the repetition of a stimulus (e.g., consumption) reduces both the perception of the stimulus and the response to that stimulus.

as capital expenditures, growth in real gross national product, and industrial production indicating that the market rationally anticipates real activity (Fama, 1981). Lee (1992) examines this question in a vector autoregressive framework that includes real stock returns and real interest rates. He finds evidence that positive, real stock return shocks appear to have a strong, positive effect on subsequent real activity. In a similar vein, Barro (1990) shows that changes in real stock market prices at a one-year lag have strong predictive information about the following year's real, nonresidential, private fixed investment and real GNP.[5]

Other empirical evidence indicates that expected excess returns vary countercyclically with current business conditions. Chen (1991) examines the expected excess returns on the value-weighted portfolio of NYSE stocks relative to recent GNP growth and expected future GNP growth and growth volatility. He finds the expected excess return is countercyclical, rising as recent output growth falls. Recent work by Pástor and Stambaugh (2001) supports this countercyclicality. They find the equity risk premium reaches its highest level during the Great Depression and the stagflationary recession of the early 1970s. Its steepest decline is likewise during the long economic expansion of the 1990s. Fama and French (1989), who show similar results, offer an interpretation based on the consumption-smoothing hypothesis (Hall, 1978; Modigliani, 1986): expected excess returns rise and fall opposite to economic prosperity because investors require a higher return during periods of low income. To induce investment in risky equity during recessions, firms have to offer a higher risk premium to entice investment, as current consumption is more highly valued. Moreover, Chen confirms the theoretical results of Abel (1988, 1999) and finds conditional expected excess returns rise as expected future GNP growth rises (to the contrary, however, he finds expected growth volatility has little explanatory power).

3. Methodology

Our approach comprises two stages of analysis. In stage 1, we estimate the risk premia associated with the Fama and French (1993) three-factor model. We select this model because of its prominence in the literature and its ability to explain returns on portfolios that display aberrations relative to the CAPM (also see Fama and French, 1996). In stage 2, we begin with a standard vector autoregressive (VAR) monetary model of the economy and augment it with three factors and their risk premia. With this model, we then test hypotheses about the dynamic interaction between investor risk premia and real macroeconomic time series such as durable goods orders, housing starts, and employment.

3.1. Estimation of the risk premia

We assume asset (or portfolio) i's return at time t may be described by the Fama and French (1993) three-factor model:

$$r_{it} = r_{0t} + \beta_{i,\mathrm{MRF}}\mathrm{MRF}_t + \beta_{i,\mathrm{SMB}}\mathrm{SMB}_t + \beta_{i,\mathrm{HML}}\mathrm{HML}_t + \varepsilon_{it}, \tag{1}$$

[5] Investors' anticipation of changes in the future *cash flows* may well drive many of these results. In contrast to these papers, we focus exclusively on the discount rate or, more precisely, the risk premium and examine how current changes in the risk premium may influence future real activity.

where r_{0t} is the return on the zero-beta portfolio (we follow Black (1972) and Lettau and Ludvigson (2001) and assume only that an asset with zero factor betas exists, not one that is risk-free). MRF_t, SMB_t, and HML_t are the factor realizations at time t; $\beta_{i,\mathrm{MRF}}$, $\beta_{i,\mathrm{SMB}}$, and $\beta_{i,\mathrm{HML}}$ are the factor sensitivities of asset i; and ε_{it} is an (idiosyncratic) error term.

Fama and French construct the MRF factor by taking the difference in returns between a (value-weighted) broad market index and a riskless rate. They construct SMB by taking the difference in returns between a portfolio of small stocks and a portfolio of big stocks each having the same weighted average book-to-market equity ratio. They construct HML by taking the difference in returns on a portfolio of high and a portfolio of low book-to-market equity portfolios each having the same weight-averaged size. Fama and French (1993) observe that MRF is "a hodgepodge of the common factors" and orthogonalize it relative to the other factors in one set of tests of their model. Similarly, we use the residuals from a regression of the market excess return on SMB and HML as MRF.

We estimate the risk premia using the Fama and MacBeth (1973) procedure. Versions of this technique very similar to ours have recently been used by Fama and French (1992), Kothari et al. (1995), Lettau and Ludvigson (2001) to test several linear asset pricing models. First, starting June 1958, we select all stocks on the Center for Research in Security Prices (CRSP) tape with at least 24 returns over the last 60 months including June and regress those returns on a constant and the three Fama–French factors.[6] The factor coefficients for each stock are saved and used as the three "betas" of the stock for the *next* 12 months. For each of these subsequent months (July through the following June), stocks are sorted into quantiles in a nested fashion based on their beta values. Following Kothari et al. (1995), we determine quantile breakpoints using only NYSE and AMEX stocks to avoid the distortional effects of the predominately small NASDAQ stocks that join the dataset in 1973. Specifically, stocks are first ranked into three groups based on their MRF betas. Within each of these groups, stocks are further ranked into three groups based on their SMB betas. Within each of these nine groups, stocks are ranked one last time into three groups based on their HML betas for a total of 27 portfolios. Returns on these portfolios are calculated as equally weighted returns of their component stocks. This process is repeated year by year ending June 1998.

Next, we calculate one set of betas for each of the 27 portfolios using the entire sample period, July 1958–June 1999. We use Eq. (1) and regress the monthly portfolio returns on a constant and the three factors. To account for the possibility of nonsynchronous trading, however, we follow Fama and French (1992) and use the Dimson (1979) correction: we include a lag of each factor in the regressions and calculate the portfolio betas by summing the two coefficients on each respective factor. This part of the analysis produces the *estimated* factor sensitivities, $\hat{\beta}_{i,\mathrm{MRF}}$, $\hat{\beta}_{i,\mathrm{SMB}}$, and $\hat{\beta}_{i,\mathrm{HML}}$, for each of the 27 portfolios.

According to the Fama–French model, the expected return on a portfolio (or any asset) may be represented by

$$E_{t-1}[r_{it}] = E_{t-1}[r_{0t}] + \gamma_{t-1,\mathrm{MRF}}\beta_{i,\mathrm{MRF}} + \gamma_{t-1,\mathrm{SMB}}\beta_{i,\mathrm{SMB}} + \gamma_{t-1,\mathrm{HML}}\beta_{i,\mathrm{HML}}, \tag{2}$$

where $\gamma_{t-1,\mathrm{MRF}}$, $\gamma_{t-1,\mathrm{SMB}}$, and $\gamma_{t-1,\mathrm{HML}}$ are the risk premia on the respective factors. We note that there are three separate risk premia in the economy, each represented by a distinct gamma; consequently, the overall risk premium for some portfolio i is $\gamma_{t-1,\mathrm{MRF}}\beta_{i,\mathrm{MRF}} +$

[6] Selection of June as the starting date is standard and is done to reduce the effects of seasonality in returns.

$\gamma_{t-1,\mathrm{SMB}}\beta_{i,\mathrm{SMB}} + \gamma_{t-1,\mathrm{HML}}\beta_{i,\mathrm{HML}}$. Investors conditionally set each gamma based on all information available *in advance of the realization of the factors and, hence, the asset's return*. We emphasize the risk premia are dynamic and forward looking; they need not be functions of current and past macroeconomic state variables.

Lastly, in the Fama–MacBeth methodology, we estimate the risk premia by cross-sectionally regressing the realized returns of the 27 portfolios on each portfolio's previously estimated factor betas month by month, July 1958–June 1999, using the following regression equation:

$$r_{it} = r_{0t} + \gamma_{t-1,\mathrm{MRF}}\hat{\beta}_{i,\mathrm{MRF}} + \gamma_{t-1,\mathrm{SMB}}\hat{\beta}_{i,\mathrm{SMB}} + \gamma_{t-1,\mathrm{HML}}\hat{\beta}_{i,\mathrm{HML}} + \varepsilon_{it}. \tag{3}$$

The coefficient for each beta is the respective factor's risk premium (plus an unanticipated factor shock[7]) for a given month. Consequently, this analysis produces a monthly time series of the *estimated* risk premia that we could denote as $\hat{\gamma}_{t-1,\mathrm{MRF}}$, $\hat{\gamma}_{t-1,\mathrm{SMB}}$, and $\hat{\gamma}_{t-1,\mathrm{HML}}$. From now on, however, we refer to the risk premia for factors MRF, SMB, and HML as RISKMRF, RISKSMB, and RISKHML, respectively.

3.2. The econometric model

In light of the findings of Patelis (1997) and Thorbecke (1997, 2000), who show a link between monetary policy and stock returns, and Elder (2001), who shows a link between federal funds rate volatility and the Treasury bill risk premium, we cast our analysis within the well-established framework of monetary policy vector autoregressions.[8] The benefit of this approach is the direct comparability of the effects of monetary policy on real macroeconomic variables with the effects of our risk premia on those same measures. Monetary models analyze the impact on the economy of changes in the cost banks incur to borrow for reserve shortfalls (i.e., the fed funds rate). Similarly, we analyze the impact on the economy of changes in the cost firms incur to borrow risky capital.

The influential papers of Bernanke and Blinder (1992), Strongin (1995) have identified two powerful measures of monetary policy. Bernanke and Blinder find that the fed funds rate responds well to its goal variables: the rate increases with inflation shocks and falls with unemployment shocks. Moreover, innovations in the federal funds rate account for meaningful variance in several measures of real activity. Strongin, on the other hand, argues that changes in nonborrowed reserves orthogonal to total reserve changes are a better measure of monetary policy. In our analysis, we include the fed funds rate and like Strongin total reserves (relative to lagged total reserves) and nonborrowed reserves

[7] A standard result of the Fama and MacBeth (1973) procedure is that the risk premium estimates are noisy and reflect the sum of each risk premium estimate and its unanticipated factor realization. (See Cochrane (2001, Chapter 12) and especially p. 243 for additional discussion and an equation that explicitly shows the unanticipated factor realizations in an excess return factor model.) In the current three-factor model, we may take the expectation of Eq. (1) and then substitute both Eq. (1) and its expectation into Eq. (2). Simplification then yields $r_{it} = r_{0t} + \beta_{i,\mathrm{MRF}}(\gamma_{t-1,\mathrm{MRF}} + s_{t,\mathrm{MRF}}) + \beta_{i,\mathrm{SMB}}(\gamma_{t-1,\mathrm{SMB}} + s_{t,\mathrm{SMB}}) + \beta_{i,\mathrm{HML}}(\gamma_{t-1,\mathrm{HML}} + s_{t,\mathrm{HML}}) + \varepsilon_{it}$ where $s_{t,\mathrm{MRF}} \equiv \mathrm{MRF}_t\text{-}\mathrm{E}_{t-1}[\mathrm{MRF}_t]$, the unanticipated factor shock for the MRF factor, and where $s_{t,\mathrm{SMB}}$ and $s_{t,\mathrm{HML}}$ are similarly defined. As a consequence, there is a potential errors-in-variables problem in subsequent analysis that uses these estimates. In later analysis, however, parameter values are not central to the conclusions of this paper. Our primary consideration is forecasting, which is not affected by parameter bias (see Davidson and MacKinnon (1993, pp. 210–211)).

[8] We thank an anonymous referee for steering us in this direction.

(relative to lagged total reserves). (See Choi and Ratti (2000) for a detailed comparison of these monetary policy indicators.) Nevertheless, Christiano et al. (1996) find an increase in inflation following positive innovations in the fed funds rate (the price puzzle). However, these authors show that if a commodity price index is included in the regression, this effect disappears. This practice has become standard and we too include a commodity price series in our VARs.

Our systems are essentially a standard monetary model augmented with asset pricing factors and risk premia. As with many common monetary models, we achieve identification by assuming a lower triangular matrix for the contemporaneous interactions among the variables (i.e., we use a Choleski decomposition). We estimate seven different systems and analyze their impulse responses. The first series in each system is *one* of the seven real macroeconomic series appearing in levels in Bernanke and Blinder (1992), namely, new durable orders, consumption, employment, housing, income, industrial production, and retail sales. The remaining variables in *each* system are inflation; a commodity price index; total reserves relative to lagged total reserves; nonborrowed reserves relative to lagged borrowed reserves; the fed funds rate; MRF, SMB, and HML; and RISKMRF, RISKSMB, and RISKHML, in this order. Table 1 describes these series, their naming conventions, and any transformations. Table 2 shows that with the exception of HOUSING, both augmented Dickey–Fuller and Phillips–Perron tests consistently find unit roots in the real macroeconomic series.[9] In addition, there is mixed evidence that INFL and FFUNDS are nonstationary.[10] In untabulated results, we also conduct Phillips–Ouliaris–Hansen tests (See Hamilton, 1994, Chapter 19) and find evidence of bivariate cointegration between inflation and FFUNDS.[11]

The Choleski decomposition of the covariance matrix plays two key roles in our impulse-response function analyses. First, by ordering the risk premia after the factors, the innovations in the risk premia are orthogonal to unanticipated shocks to their respective factors. Consequently, this decomposition eliminates the need to extract the factor shocks from the risk premia estimates and allows a direct inference with regard to the impact of innovations in the premia themselves.[12] Second, as in Strongin (1995), placing NBRX after TRX allows the innovation in NBRX to capture the change in nonborrowed reserves orthogonal to TRX, Strongin's monetary policy measure.

We estimate all systems as vector autoregressions in levels:

$$\mathbf{y}_t = \boldsymbol{\mu} + \sum_{i=1}^{k} \boldsymbol{\Gamma}_i \mathbf{y}_{t-i} + \boldsymbol{\varepsilon}_t. \tag{4}$$

$\mathbf{y}_t$ is a p-vector of one real macroeconomic variable, the monetary policy variables, the risk factors, and their respective premia. $\boldsymbol{\mu}$ is a vector of drift constants (we let $\hat{\boldsymbol{\mu}}$ represent the *estimated* vector), and $\boldsymbol{\Gamma}_i$ is the $p \times p$ matrix of coefficients at lag i (we let $\hat{\boldsymbol{\Gamma}}_i$ represent the

[9] New econometric evidence in Elder and Kennedy (2001) indicates the Dickey–Fuller t-test dominates the DF F-test for series that do not trend.

[10] Prices are frequently found to be integrated of order two, and consequently, we use inflation rather than prices, as does Thorbecke (1997). In addition, we test whether PCOM is integrated of order two (I(2)). We check this using a Dickey and Pantula (1987) test and reject an I(2) process in favor of a unit-root-with-drift process.

[11] We have also checked for the robustness of our results to the specification of cointegration in an error-correction model (Johansen, 1995). We find the results are qualitatively unchanged.

[12] See note 7.

Table 1
Description of financial and macroeconomic data series

Series name	Data source	Description of series	Transformations	Period
IPROD	DRI: IP	Industrial production: total (1992 = 100, SA)	Log level	1959:1–1998:9
CONSUME	DRI: GMCQ	Personal consumption expenditures (1992 $, SAAR)	Log level	1959:1–1998:9
EMPLOY	DRI: LHNAG	Nonagriculture civilian labor (thousands, SA)	Log level	1959:1–1998:9
HOUSING	HSMW, HSNE, HSSOU, HSWST	Housing starts by region (thousands of units)	Log level of sum	1959:1–1998:9
INCOME	DRI:GMPYQ	Personal income (1992 $, SAAR)	Log level	1959:1–1998:9
NEWDUR	DRI: MDOQ	New orders of durables (1992 $)	Log level	1959:1–1998:9
RETAIL	DRI: RTRR	Retail sales: total (million $)	Log of CPI-deflated level	1959:1–1998:9
INFLATION	DRI: PUNEW	CPI – Urban (1982–1984 = 100, SA)	Log difference	1959:1–1998:9
PCOM	DRI: PSCCOM	Spot market price index of all commodities (1967 = 100, NSA)	Log level	1959:1–1998:9
TRX	DRI: FMRRA	Total Reserves Adjusted for required reserves changes (million $, SA)	Ratio to lagged level	1959:1–1998:9
NBRX	DRI: FMRNBC	Nonborrowed reserves + extended credit adj. for required reserve changes (million $, SA)	Ratio to lagged level of total reserves	1959:1–1998:9
FFUNDS	DRI: FYFF	Federal funds interest rate (% per annum, NSA)	None	1959:1–1998:9
MRF	Kenneth French	Excess return of the market portfolio over the riskfree of Fama and French (1993)	Orthogonalized to SMB and HML	1953:7–1999:6
SMB	Kenneth French	Average return on the three smallest benchmark portfolios of Fama and French (1993) less the average return on the three largest	None	1953:7–1999:6
HML	Kenneth French	Average return on the two benchmark value portfolios of Fama and French (1993) less the average return on the two growth portfolios	None	1953:7–1999:6

Table 1 (*continued*)

Series name	Data source	Description of series	Transformations	Period
RISKMRF	CRSP common stock returns	Risk premia on MRF calculated using Fama–MacBeth procedure on stock returns	None	1953:7–1999:6
RISKSMB	CRSP common stock returns	Risk premia on SMB calculated using Fama–MacBeth procedure on stock returns	None	1953:7–1999:6
RISKHML	CRSP common stock returns	Risk premia on HML calculated using Fama–MacBeth procedure on stock returns	None	1953:7–1999:6

Notes: The table describes the monthly financial and macroeconomic data series used in the systems analyzed. The asset returns used to calculate the risk premia are from the Center for Research in Security Prices (CRSP). All macroeconomic series are seasonally adjusted series from the DRI Database. NSA = not seasonally adjusted. SA = seasonally adjusted. SAAR = seasonally adjusted at an annual rate.

Table 2
Tests for nonstationarity

Series	Augmented Dickey–Fuller tests		Phillips–Perron tests	
	t	Z_{DF}	Z_ρ	Z_t
IPROD	−2.41	−10.57	−8.89	−2.40
CONSUM	−1.88	−6.58	−4.78	−1.85
EMPLOY	−2.00	−10.23	−3.52	−1.12
HOUSING	−3.36	**−32.60**	**−29.85**	−3.95
INCOME	−2.01	−6.35	−3.22	−1.62
NEWDUR	−2.81	−17.66	−17.77	−3.01
RETAIL	−2.27	−10.15	−11.33	−2.48
INFL	−2.17	−10.25	**−156.18**	**−9.59**
PCOM	−2.84	−25.02	−3.89	−1.19
TRX	**−3.57**	**−37.30**	**−483.18**	**−18.61**
NBRX	**−4.82**	**−46.27**	**−53.14**	**−5.36**
FFUNDS	−2.52	**−15.27**	−12.23	−2.56
MRF	**−22.75**	**−497.71**	**−513.38**	**−22.67**
SMB	**−18.34**	**−395.20**	**−400.75**	**−18.37**
HML	**−18.61**	**−403.66**	**−392.91**	**−18.54**
RISKMRF	**−20.54**	**−449.05**	**−408.99**	**−20.51**
RISKSMB	**−4.30**	**−81.63**	**−369.38**	**−17.65**
RISKHML	**−4.64**	**−96.35**	**−387.42**	**−18.25**

Notes: The table describes both augmented Dickey–Fuller (DF) and Phillips–Perron (PP) tests for all series used in the VAR systems 1959:2–1998:9. For the DF tests, lag length is determined using the Akaike criterion (AIC). For PP statistics, the Newey-West estimator with a lag truncation parameter of six is used to estimate λ^2. We include a constant in the regressions of nontrending series (INFL, TRX, NBRX, FFUNDS, MRF, SMB, HML, RISKMRF, RISKSMB, and RISKHML). For the other (trending) series we include both a constant and trend. Values in bold denote an observed significance level that exceeds 1%. See Hamilton (1994, Chapter 17) for further details.

estimated $p \times p$ matrix at lag *i*). $\boldsymbol{\varepsilon}_t \sim \mathrm{nid}(\mathbf{0},\boldsymbol{\Sigma})$ is the vector of innovations. For each system, we specify the number of lags based on likelihood ratio tests, and in each case, these tests select a lag length of 12. We plot impulse-response functions using the moving average representation of (4) and include 95%-confidence bands that we generate using Monte Carlo techniques and 0.025 and 0.975 fractiles, as recommended by Sims and Zha (1999).

4. Fama–French risk premia and equilibrium asset pricing

This section explores the relationship between the dynamics of the risk premia derived from the empirical Fama–French model with the dynamics of the premia in the general equilibrium models described in Section 2. These theoretical models link asset pricing to consumption and production and hence *provide the essential theoretical link between macroeconomic fluctuations and the premia for risk we study*. The majority of these models imply that the risk premium should vary positively with future expected consumption and future consumption uncertainty. We now compare the characteristics of our measures of risk compensation with those of theory.[13]

Chen (1991) and Kandel and Stambaugh (1990), whose approach we adopt, compare excess returns to consumption expectations. Their insight is to use state variables such as the differences in yields on bonds of different risk or maturity as instruments to extract investors' expectations of consumption and consumption volatility. With these extracted expectations in hand, these authors then analyze the relationship between excess returns and expectations of consumption and consumption volatility. Because excess returns are the noisy realizations of (i.e., a linear function of) the risk premia determined ex ante, we undertake a parallel analysis and link the risk premia we study to the same extracted consumption expectations.

Several variables sensitive to macroeconomic fluctuations have been proposed as measures of economic conditions or proxies for the information set of investors. With respect to intertemporal asset-pricing, the value of any such proxy is its ability to forecast changes in the consumption–investment opportunity set. If correctly chosen, these measures should capture information important to investors as they set asset prices and the premia for various sources of risk. Fama and French (1989) find three variables that predict stock and bond returns in excess of the riskless rate: the default spread between yields on high- and low-grade corporate bonds, the term spread between safe long- and short-term bonds, and the dividend yield on the market portfolio.

Chen (1991) and Kandel and Stambaugh (1990) also study excess market returns but cast their tests within an asset pricing framework that links their analysis directly to consumption. Their analysis proceeds in two steps. First, they fit (condition) consumption growth levels and volatility to current economic state variables. Second, they regress realized excess returns on those fitted values. In the second regression, the Chen–Kandel–

[13] It may seem peculiar to compare a nonconsumption-based model's risk premia to consumption. However, as Cochrane (2001, p. 170) writes: "…derivations make clear that CAPM and ICAPM [Fama and French (1993) view their model an ICAPM model] are not alternatives to the consumption-based model; they are special cases of that model…We just added assumptions that allowed us to substitute other variables in place of c_t [consumption]." The value of factors in multifactor models derives from the factors' links to consumption. The empirically superior performance of nonconsumption factor models is due to the greater precision with which factors can be measured as opposed to consumption.

Stambaugh technique directly relates the dynamics of the risk premium to conditional consumption levels and volatility and tests several implications of the models described in Section 2, such as Abel (1988, 1999). Specifically, this second regression investigates whether the risk premium rises as expected future consumption rises and as expected future consumption volatility increases. More broadly, this regression provides insight into whether the risk premium varies countercyclically with the business cycle, as predicted by many theoretical asset pricing models.

We apply the Chen–Kandel–Stambaugh technique to a system of equations estimated by the generalized method of moments to control for a generated regressors problem.[14] For expositional clarity, however, we discuss the approach as if it proceeded sequentially. We estimate expected future consumption growth (the first difference of CONSUME) as a linear function of current state variables (see Table 3, Eq. (i)). Our instruments are very similar to those in Chen (1991) and Kandel and Stambaugh (1990): They include the average yield spread between Moody's Baa- and Aaa-rated corporate bonds (DEF), the yield spread between the 10-year constant maturity US Treasury bond and the one-month Treasury bill annualized (TERM),[15] the dividend yield on the S&P 500 Composite Index (DIVYLD), and the nominal one-month T-bill rate annualized. All data are from the DRI Database except the one-month T-bill rate, which comes from Ibbottson Associates Inc. The fitted values then estimate expected future consumption growth. In panels A and B of Table 3, (i) shows these results for three- and six-month horizons.

To estimate expected future consumption volatility, we project the absolute value of the residuals from Eq. (i) times $\sqrt{\pi/2}$ on the same instruments. These fitted values then approximate the standard deviation of conditional future consumption volatility. In panels A and B of Table 3, (ii) shows these results for three- and six-month horizons.

We next relate these estimates of the market's expectation of future consumption growth and volatility to our estimated Fama–French risk premia. We project the risk premia on the fitted consumption measures from Eqs. (i) and (ii) in Table 3 for both three- and six-month horizons. To account for the effect of habit formation, we also include realized current consumption along with five of its lags. As the estimation results for (iii) and (iv) show in panels A and B of Table 3, the signs on expected consumption growth and volatility are generally positive for RISKMRF and RISKSMB at both horizons; though they are not always significant. This agrees with the predictions of theory: as expected consumption or its volatility increases, the risk premia rise.[16] To the contrary, we observe a negative relationship between RISKHML and the consumption measures in Eq. (v). However, these coefficients are never significant. In addition, when we project excess returns from the CRSP value-weighted market return on future expected consumption growth and volatility, we find a significant and positive relationship at both the three- and six-month horizon as shown in Eq. (v′).

The graphs of these risk premia in Fig. 1 also add support to these inferences. At the trough of recessions or shortly thereafter, recent consumption is very low and future expected consumption and consumption volatility are very high. As expected, the graphs

[14] We thank an anonymous referee for bringing this issue to our attention.

[15] Abel (1999) shows that the equity premium comprises a term premium and a risk premium. Because long-term bond prices are more sensitive to interest rate volatility than short-term bonds, investors demand compensation for this source of risk in addition to the cash flow risk from stocks.

[16] See note 3.

Table 3
Expected consumption means and volatility

$$c_{t+\tau} = \alpha + \beta_1 \mathrm{DEF}_t + \beta_2 \mathrm{TERM}_t + \beta_3 \mathrm{DIVYLD}_t + \beta_4 \mathrm{TB}_t + \varepsilon_{t+\tau} \quad \text{(i)}$$

$$\sqrt{\pi/2}|\hat{\varepsilon}_{t+\tau}| = \alpha + \beta_1 \mathrm{DEF}_t + \beta_2 \mathrm{TERM}_t + \beta_3 \mathrm{DIVYLD}_t + \beta_4 \mathrm{TB}_t + \eta_{t+\tau} \quad \text{(ii)}$$

$$\mathrm{RISKMRF}_t = \alpha + \beta_1 \hat{\mu}_{c,t+\tau} + \beta_2 \hat{\sigma}_{c,t+\tau} + \textstyle\sum_{j=0}^{5} \delta_j c_{t-j,t-j-1} + e_{t,\mathrm{MRF}} \quad \text{(iii)}$$

$$\mathrm{RISKSMB}_t = \alpha + \beta_1 \hat{\mu}_{c,t+\tau} + \beta_2 \hat{\sigma}_{c,t+\tau} + \textstyle\sum_{j=0}^{5} \delta_j c_{t-j,t-j-1} + e_{t,\mathrm{SMB}} \quad \text{(iv)}$$

$$\mathrm{RISKHML}_t = \alpha + \beta_1 \hat{\mu}_{c,t+\tau} + \beta_2 \hat{\sigma}_{c,t+\tau} + \textstyle\sum_{j=0}^{5} \delta_j c_{t-j,t-j-1} + e_{t,\mathrm{HML}} \quad \text{(v)}$$

$$r_{e,t+\tau} = \alpha + \beta_1 \hat{\mu}_{c,t+\tau} + \beta_2 \hat{\sigma}_{c,t+\tau} + \textstyle\sum_{j=0}^{5} \delta_j c_{t-j,t-j-1} + e_t \quad \text{(v}'\text{)}$$

Equation	α	β_1	β_2	β_3	β_4
Panel A: $\tau = 3$ months					
(i)	0.0122	0.6274	−0.0117	−0.0779	−0.1348
	(0.0000)	(0.0002)	(0.7661)	(0.3500)	(0.0001)
(ii)	0.0054	−0.0727	−0.0627	0.1788	−0.0463
	(0.0006)	(0.5378)	(0.0213)	(0.0060)	(0.0483)
(iii)	−0.0049	2.2861	−0.9567		
	(0.8315)	(0.0418)	(0.7117)		
(iv)	−0.0304	0.9575	3.1674		
	(0.0126)	(0.1228)	(0.0261)		
(v)	0.0141	−0.9895	−0.8688		
	(0.3073)	(0.1508)	(0.5567)		
(v′)	−0.2568	23.2278	27.0087		
	(0.0980)	(0.0020)	(0.1159)		
Panel B: $\tau = 6$ months					
(i)	0.0211	1.2347	−0.0516	−0.0537	−0.2861
	(0.0000)	(0.0000)	(0.4466)	(0.6948)	(0.0000)
(ii)	0.0058	−0.0354	−0.0883	0.1841	−0.0273
	(0.0277)	(0.8264)	(0.0476)	(0.0637)	(0.5608)
(iii)	−0.0051	1.0224	−0.5702		
	(0.8325)	(0.0929)	(0.7695)		
(iv)	−0.0362	0.8405	2.7087		
	(0.0242)	(0.0296)	(0.0489)		
(v)	0.0121	−0.4979	−0.5861		
	(0.3965)	(0.1695)	(0.6340)		
(v′)	−0.2016	11.0643	16.5428		
	(0.1943)	(0.0009)	(0.2043)		

Notes: $c_{t+\tau,t} = \mathrm{CONSUME}_{t+\tau} - \mathrm{CONSUME}_t$. DEF_t = yield spread between the Moody's Baa- and Aaa-rated corporate bond portfolios at time t. TERM_t = yield spread between the 10-year constant maturity US Treasury bond and the annualized, continuously compounded, nominal one-month Treasury bill. DIVYLD_t = dividend yield on the S&P 500 Composite Index, and TB_t = the annualized, continuously compounded, nominal one-month T-bill yield of month t. These data are from the DRI Database except the one-month T-bill rate, which comes from Ibbottson Associates Inc. $\hat{\mu}_{c,t+\tau}$ and $\hat{\sigma}_{ct+\tau}$ are estimated conditional expected consumption growth means and standard deviations for period t to $t+\tau$ using the fitted values of Eqs. (i) and (ii), respectively. $r_{e,t+\tau}$ is the annualized, continuously compounded, monthly return of the CRSP value-weighted portfolio in excess of TB_t. The coefficients are estimated in two systems, one with Eqs. (i)–(v) and a second with (i)–(iv) and (v′); only the coefficients for Eq. (v′) are reported from the second system. Estimation is through the generalized method of moments (GMM) for the period 1959:7–1998:9–τ. The Newey and West (1987) kernel with a window of $\tau - 1$ lags/leads is used as the weighting matrix to correct for heteroskedasticity and serial correlation; the exogenous

show the risk premia are often large at the end of recession or very early into the subsequent expansion.

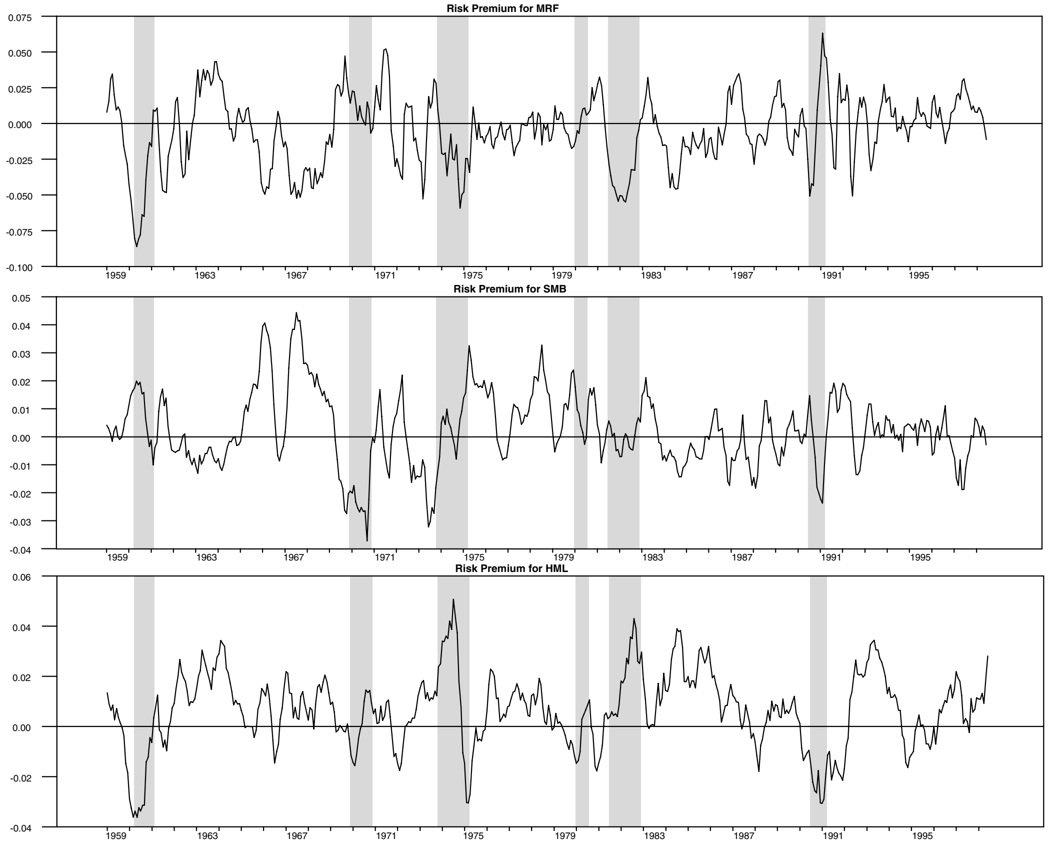

Fig. 1. Time series of risk premia (*Notes*: This figure graphs the time series of the risk premia for MRF, SMB, and HML over the period 1959:1 to 1998:9 with shaded regions representing recessions as determined by the NBER. The unanticipated factor shocks (see note 7) have been subtracted, and the risk premia have been smoothed by using a moving average of each risk premium over a window of three lags and three leads. The unanticipated factor shocks are approximated as the difference between each respective factor's current realization and its six-lag moving average. (This approach is used only in this figure for the purposes of graphing, not in any statistical analyses.))

Although they pose many future research questions, these initial results are generally consistent with the economic theory and intuition outlined earlier in Section 2. Risk premia appear to be countercyclical barometers of rationally anticipated economic conditions and reflect a consumption-risk mechanism.

5. Impulse-response analysis

We are interested in how investors' aversion to risk factors leads future real activity. As we noted earlier, finding the stock market forecasts real activity is not surprising under the assumption of rational markets. The stock market value is the present value of future profit distributions, which intuitively should rise and fall with economic prosperity. This thinking focuses on the numerators or cash flows in the discounted dividend model implied by Lucas (1978). In contrast, the focus of this paper is on the denominators or the discount rates.

We test the effect of changing risk premia in the discount rate using impulse-response techniques. We test whether our risk premia "cause" changes in the economy, examine the direction and magnitude of these changes, and compare them to responses to shocks in monetary policy variables. It is important to note that as in any nonexperimental inquiry, true causality may never be tested: an economic factor exogenous to the system may truly drive the results and its correlation with a variable within the estimated system may lead to a spurious causal inference. The hope is simply that enough other time series related to the variable(s) of interest are included such that valid inferences about the potentially causal variables are possible.[17] By casting our analysis within the well-established VARs of monetary policy research, we believe we have met this requirement.

We turn now to the central issue of this research: the responses of real economic variables to impulses in the risk premia. In this discussion, we emphasize new durable goods orders because this series measures future demand for production and employment. Durable goods include capital equipment such as industrial machinery, computers, electrical control instruments, trucks, aircraft, and ships, and new orders of durables often precede new investment in plant and equipment. Consequently, a rise in new orders for durable goods indicates firms expect a subsequent rise in demand for their products. As a result, NEWDUR should be a strong measure of real economic activity.

Fig. 2 shows that NEWDUR is very quick to respond to a one-standard deviation increase in RISKSMB inducing a −0.75% drop in new durable orders at the two-month horizon. Moreover, the effects persist. In general, new durable orders are lower by about −0.50% over the 3–12-month horizon. The subsequent impulse-response pattern is similar to that of the standard VAR with stationary time series attenuating to almost no effect within 18 months. Somewhat surprising, at least on the surface, is the lack of any significant effect of innovations in the risk premia on the MRF and HML factors. This issue is discussed in detail in Section 6 below.

The factors MRF, SMB, and HML themselves are also part of each system. The primary reason for their inclusion is to isolate the effect of the risk premia net of factor shocks (see note 7 above) by way of the Choleski decomposition. Nonetheless, in an ICAPM framework, the state variables must be informative regarding future consumption–investment

[17] We thank an anonymous referee for pointing this out to us.

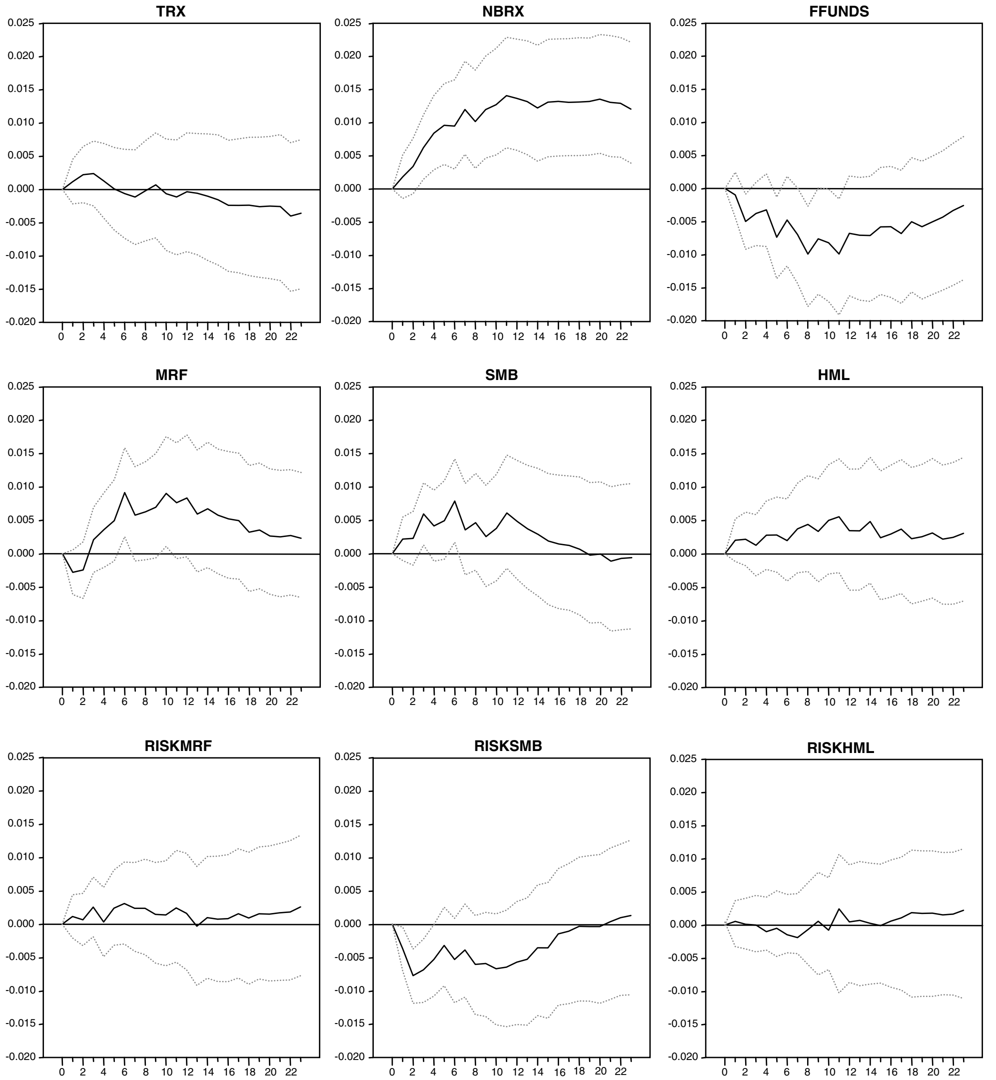

Fig. 2. Responses of NEWDUR to financial shocks (*Notes*: Figures show the dynamic responses of NEWDUR to one-standard deviation innovations in TRX, NBRX, FFUNDS, MRF, SMB, HML, RISKMRF, RISKSMB, RISKHML over a 24-month time horizon. The system is lower triangular and consists of NEWDUR, INFL, PCOM, TRX, NBRX, FFUNDS, MRF, SMB, HML, RISKMRF, RISKSMB, and RISKHML arranged in this order and all with 12 lags. The solid lines represent the impulse-response functions; the upper and lower dotted lines indicate the 95%-confidence-interval boundaries estimated using Monte Carlo techniques. The real variables are in logarithmic levels; hence, the response functions represent percentage changes (in decimal form)).

opportunities; therefore, we present the response functions of new durable orders to innovations in each of the three factors in Fig. 2. We observe innovations in MRF and SMB induce significant and expansionary effects on new durable orders ranging from 0.25% to 0.90% over the 3- to 12-month horizon. Consequently, an asset that positively covaries with these factors should earn a premium for it does better during booms, when firms' new

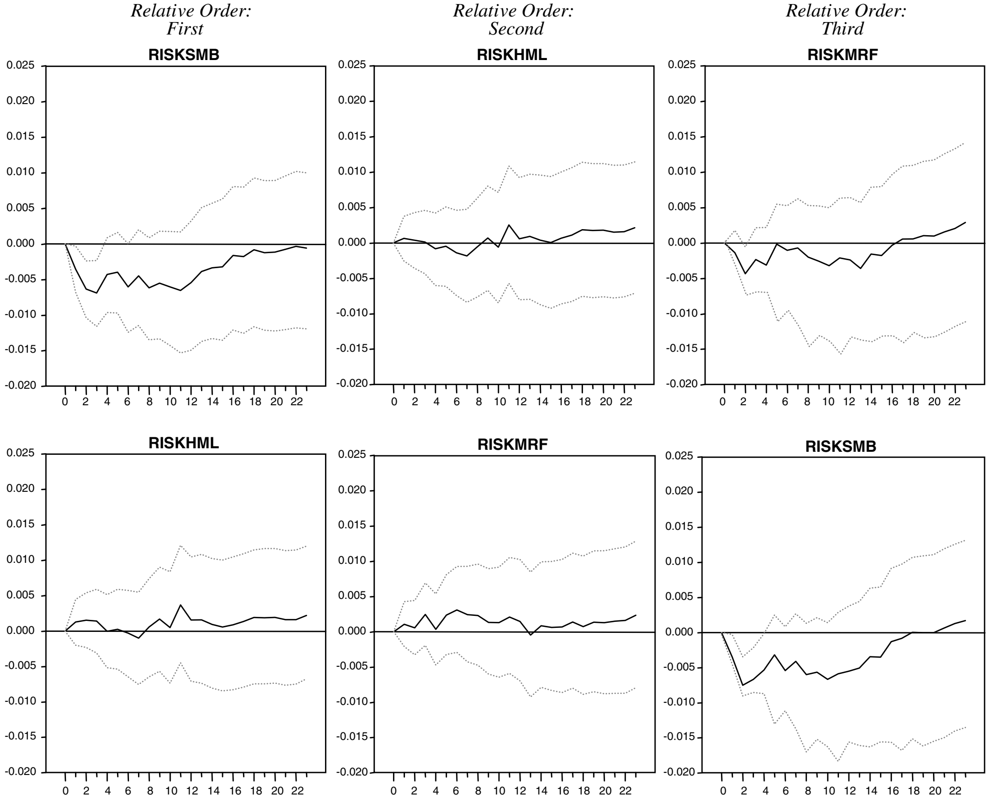

Fig. 3. Responses of NEWDUR to shocks in reordered risk premia (*Notes*: Figures show the dynamic responses of NEWDUR to one-standard deviation innovations in RISKMRF, RISKSMB, RISKHML over a 24-month time horizon for two re-orderings of the risk premia as compared to Fig. 2. The column heading indicates the relative order. Except for the reordering of the risk premia, the systems here are identical to the system in Fig. 2. See the notes for Fig. 2 for additional details).

durable goods orders and productive capabilities are increasing, and worse during busts, when these orders and capabilities are decreasing–clearly undesirable characteristics from the vantage point of a risk-averse investor.

It is well known that the order of the variables matters in a recursive VAR. To have an effect, an innovation in one time series must have a component orthogonal to innovations in all variables that precede it. Although the risk factors are essentially mutually orthogonal (as described above, we orthogonalize MRF relative to SMB and HML; moreover, Fama and French (1993) find SMB and HML have a correlation of only −0.08), their risk premia need not be. We therefore include the analysis of Fig. 3, which reorders the risk premia relative to one another but still places them after all other variables in each system.[18]

[18] All six possible orders of the risk premia were examined, but the impulse-response patterns did not materially differ from those reported in Figs. 2 and 3.

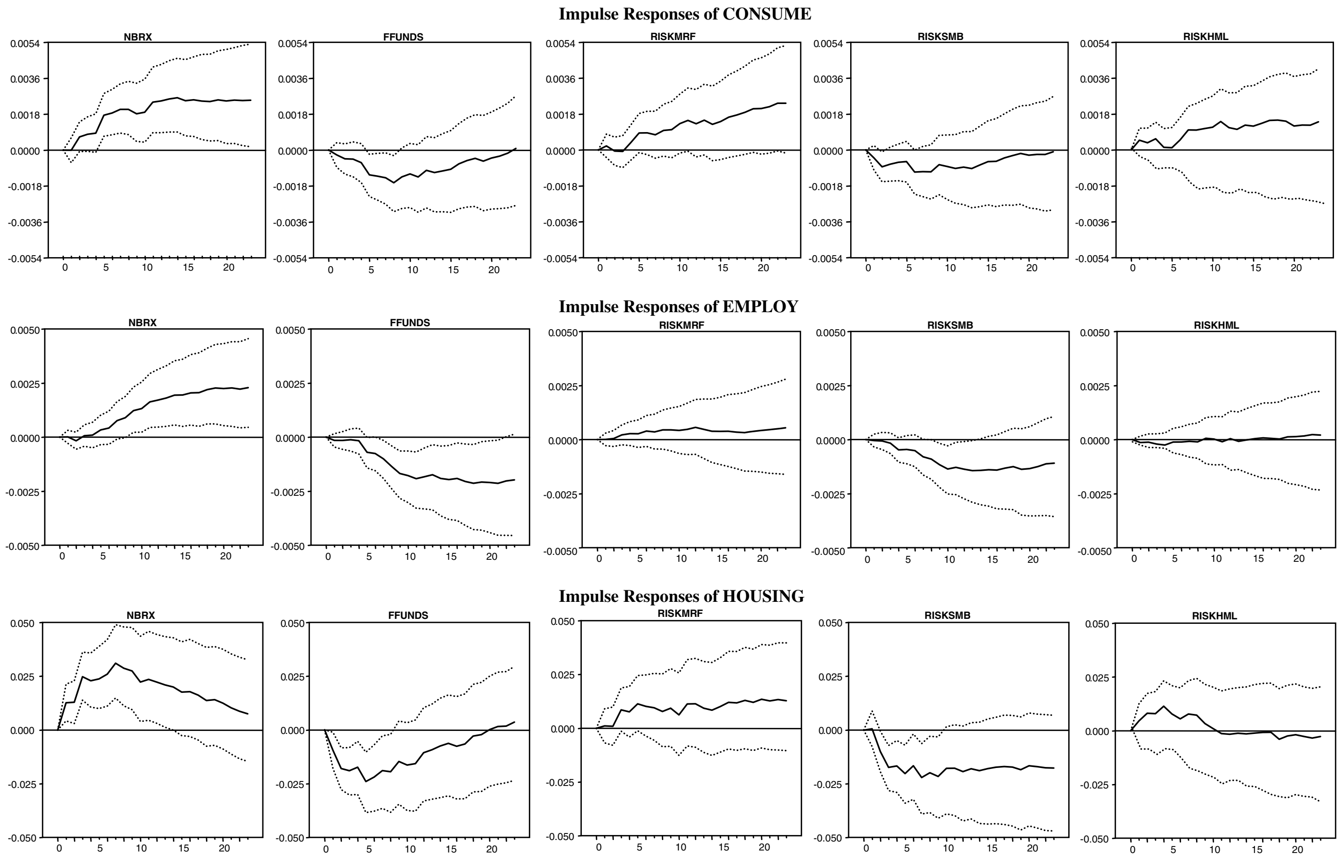

Fig. 4. Responses of real activity measures to risk premia shocks (*Notes*: Figures show the dynamic responses of the real macroeconomic variables CONSUME, EMPLOY, HOUSING, INCOME, IPROD, and RETAIL to one-standard deviation innovations in NBRX, FFUNDS, RISKMRF, RISKSMB, RISKHML over a 24-month time horizon in *separate* systems. These real macrovariables replace IPROD in the system of Fig. 2. See the notes for Fig. 2 for additional details).

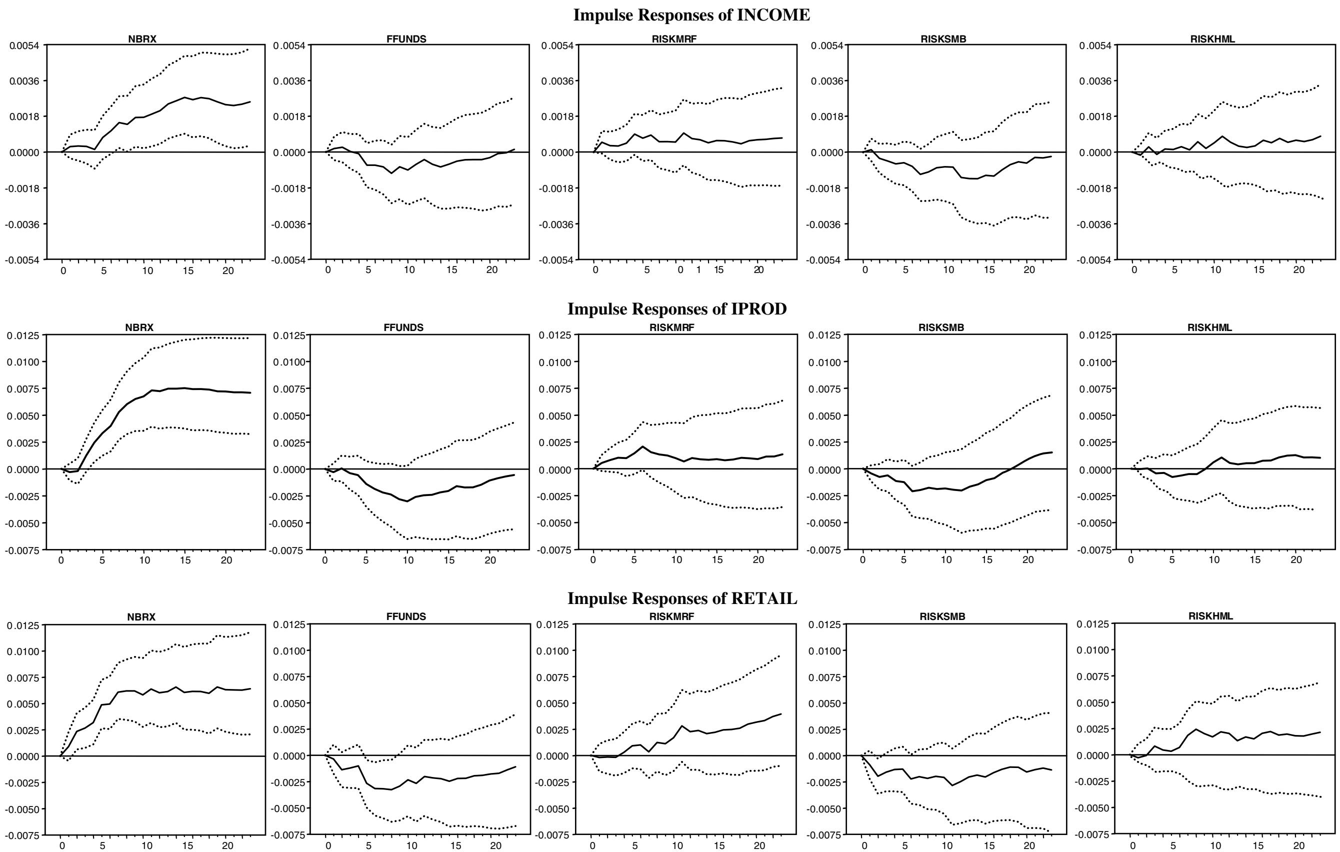

Fig. 4 (*continued*)

We see in Figs. 2 and 3 that despite ordering the risk premia after inflation, all the monetary variables, and the factors, the risk premium on SMB remains consistently informative about future states of the economy. In general, a standard deviation shock to RISKSMB induces about a −0.50% to −0.75% drop in NEWDUR over a 2- to 12-month horizon. These results hold whether RISKSMB is ordered prior to the other risk premia, as in the first system of Fig. 3, or subsequent to the other risk premia, as in the second system of Fig. 3. Moreover, these responses have magnitudes and durations very similar to those from shocks to FFUNDS (in Fig. 2) despite the ordering of the risk premia subsequent to this monetary policy variable. The general lack of informativeness of RISKMRF (RISKHML) persists, however, even when ordered first among the risk premia in Fig. 2 (Fig. 3).

Fig. 4 shows the responses of several other real variables to innovations in the risk premia. In each case, we observe similar magnitudes of responses to FFUNDS and RISKSMB shocks. As with the responses of new durable orders in Figs. 2 and 3, RISKMRF and RISKHML play little role. HOUSING responds to a one-standard deviation innovation in RISKSMB by −2.0% to −2.5%, an effect that persists over a 4- to 24-month horizon. In addition, we observe the response of HOUSING to RISKSMB is substantially more persistent than the response of this series to FFUNDS. This high level of responsiveness is logical in light of the sensitivity of home construction to mortgage rates and, in turn, the central role the risk premium plays in this borrowing rate.

Fig. 4 also shows EMPLOY falls significantly in response to a one-standard deviation shock to RISKSMB. Similar to its response to FFUNDS innovations, EMPLOY decreases by about 0.15% following a one-standard deviation shock to RISKSMB, a reduction that persists over a 10- to 24-month horizon. Consumption and income are about half as responsive as EMPLOY to this shock though these series' responses are rarely significant.

The real responses commonly studied in monetary policy research are those of industrial production. Strongin (1995), who argues convincingly for NBRX as the best indicator of monetary policy, finds that a one standard deviation shock to this measure in a five-variable VAR results in a positive and significant 0.80–0.85% response in industrial production at the 12- to 24-month horizon. In Fig. 4 we see a very similar response in IPROD. Somewhat different from Strongin's results is the observation that FFUNDS retains a significant—though much smaller—impact, despite its following NBRX in the causal ordering. A one-standard deviation shock to FFUNDS leads to more than a −0.25% drop in industrial production over an 8- to 12-month horizon. Another general observation is that the responses to NBRX appear more persistent than those of FFUNDS. Lastly, retail sales respond similarly to industrial production in magnitude following a one-standard deviation innovation in RISKSMB.

6. Discussion

The lack of significance of the risk premium on MRF may reflect the perennial lack of significance of this factor's premium in Fama–MacBeth tests of the CAPM that include some measure of size (see, e.g., Banz, 1981; Fama and French, 1992). In accordance with this research, we find that shocks to the risk premium for size are meaningful and contractionary. This result is also consistent with Hardouvelis and Wizman (1992), who study the risk premium on size over the business cycle using the firm-size augmented CAPM.

Moreover, we observe similarities between our capital-budgeting story, the strength of the response functions to innovations in the small firm premium, and mechanisms of monetary policy transmission. Two such mechanisms seem relevant: the balance sheet of borrowers and the credit channel (see, e.g., Bernanke, 1993; Gertler and Gilchrist, 1994).

In both views of monetary policy, restrictions on the access to new capital are central. When the Fed reduces money supply or raises the fed funds rate, in addition to raising interest rates and directly curbing economic expansion and inflation, the consequent higher cost of capital impairs the collateral value of firm balance sheets, reduces their ability to borrow at favorable rates, and lowers their capital expenditures. Balance sheet effects amplify the actions of the Fed. The firms most dependent on their balance sheets to attain credit—typically through bank loans—are small firms, the very firms that load most heavily on SMB.

The mechanism of the credit channel of monetary policy transmission differs: it is the unwillingness of banks to lend that is at issue not the lack of creditworthiness of borrowers. Following a tightening of monetary policy, banks reduce the number of new loans they make and invest more in securities as they rebalance their loan portfolios to meet reserve requirements. Again, as bank loans are made disproportionately to small firms—firms with limited access to the capital markets—the effect of tighter monetary policy is propagated through the denial of new credit to small firms.

Like the economics underlying these mechanisms of monetary policy transmission and the work of Hardouvelis and Wizman, our results imply that an increase in the risk premium on the SMB factor raises the cost of capital to small firms and induces contractions in the economy. It is the small and vulnerable firms that induce macroeconomic fluctuations, not the established firms that constitute the lion's share of financial market capitalization:

> An economy in which economic hardships are not equally distributed across firms is more vulnerable to adversity. A mild recession may turn into a severe recession following a wave of bankruptcies by marginal firms unable to refinance their obligations at relatively low costs (Hardouvelis and Wizman, 1992).

We observed in the last section that innovations in the risk premium on HML typically have an insignificant impact on the real variables in Figs. 2–4. This lack of significance becomes less surprising, however, when the results of Liew and Vassalou (2000) are considered. The Liew–Vassalou paper examines the ability of the Fama–French factors to forecast GDP growth in 10 countries. Though the factors are statistically meaningful in most countries, in the US there is a fairly consistent result: SMB is statistically significant as a forecaster of GDP but HML is not (see, e.g., Table 9 in Liew and Vassalou, 2000). If the state variable does not contain information about future consumption–investment opportunities, it is not priced in an ICAPM framework; its risk premium must be zero. As we use US data, this evidence provides a plausible explanation for the lack of significance of the HML premium in our impulse-response functions.

7. Conclusion

This paper provides new insight into the interrelationship between the financial markets and real economic conditions. Basic finance teaches corporate managers to assess the profitability of new projects by comparing the return on these new investments to a risk-adjusted cost of capital. This discount rate includes the premium demanded by the firm's shareholders for bearing the risk of that stock. A rise in the risk premium of investors may therefore quash previously feasible projects. This reduction in investment may in turn lead to lower demand for new durable goods and labor. A decline in income, new housing starts, consumption, and retail sales may then follow.

We show that innovations in the risk premia on the Fama and French (1993) three-factor asset pricing model have significant real effects similar to innovations in measures of monetary policy such as the fed funds rate. Specifically, we find that positive shocks to the risk premium to which small firms are most sensitive, the premium on the SMB factor, predict significant and persistent declines in new durable orders, employment, housing starts, and industrial production. We infer that a rise in this risk premium affects the real economy through its impact on the investment decisions of small firms similar to the credit channel mechanism of monetary policy transmission.

It is tempting to view innovations in risk premia as playing a truly causal role in the behavior of the real series. As with any nonexperimental inquiry, however, the question of true causality remains an open one. Rejection of noncausality implies that the dynamic behavior of one time series *forecasts* later dynamic behavior of a second. Whether one series actually *causes* the other is indeterminate. We stress again that we have not shown true causal links. Our results are simply consistent with this interpretation.

As a topic of future research, we propose that dynamic risk premia may act as important financial mechanisms for business cycle propagation, and further, that risk premia may be new measures of optimism that act as leading indicators of recessions and recoveries. Additionally, risk premia might act as sensitive barometers of investors' expectations of future prosperity and complement existing indicators of consumer confidence. It is our hope that much new research follows to explore these interesting possibilities.

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