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How Much Do Investors Care About Macroeconomic Risk? Evidence from Scheduled Economic Announcements

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Abstract

Stock market average returns and Sharpe ratios are significantly higher on days when important macroeconomic news about inflation, unemployment, or interest rates is scheduled for announcement. The average announcement-day excess return from 1958 to 2009 is 11.4 basis points (bp) versus 1.1 bp for all the other days, suggesting that over 60% of the cumulative annual equity risk premium is earned on announcement days. The Sharpe ratio is 10 times higher. In contrast, the risk-free rate is detectably lower on announcement days, consistent with a precautionary saving motive. Our results demonstrate a trade-off between macroeconomic risk and asset returns, and provide an estimate of the premium investors demand to bear this risk.

I. Introduction

The link between macroeconomic risk and security returns is central to financial economics. While a lot of relevant information about the economy arrives randomly over time, certain important macroeconomic news is released in the form of prescheduled announcements, whose dates are known months in advance. Investors do not know what the news will be, but they do know that

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there will be news. If asset prices respond to this news, the risk associated with holding the affected securities will be higher around announcements, and this will be anticipated by rational investors.

An extensive prior literature, which we discuss below, presents evidence consistent with a higher conditional risk of holding financial assets ahead of macroeconomic announcements. Risk-averse investors who know that they will be exposed to higher risk should demand, and in equilibrium receive, a higher expected excess return during those times (provided that this risk is priced by investors as a group). Stock returns should thus be predictably higher on announcement days, and this is the main hypothesis we explore in this paper (rather than studying the impact of announcement surprises on realized returns). For example, if investors prefer to avoid inflation risk, then times of inflation announcements must be times of higher average excess returns over a sufficiently long time period (one in which the average surprise equals 0).

Consistent with this general idea, we find that average U.S. stock market returns from 1958 to 2009 are significantly higher on days when important macroeconomic news is scheduled to be announced. On days when the consumer price index (CPI), producer price index (PPI), employment figures, or Federal Open Market Committee (FOMC) decisions are released, excess market returns average 11.4 basis points (bp) versus only 1.1 bp for all the other days. These figures imply that compensation for bearing macroeconomic announcement risk accounts for a large portion of the equity risk premium, as over 60% of the cumulative annual excess return is earned on just 13% of the trading days, whose timing is known to investors well in advance. Conversely, the risk premium for holding stocks at other times is very low, with the average excess return on those days not being statistically distinguishable from 0.[1]

Despite this evidence of a significantly higher announcement-day risk premium, we find that the realized volatility of daily stock market returns is only 4% higher. The effect on implied volatility is larger than for realized volatility, but the magnitudes are still much lower than those for the difference in returns. The stock market's realized Sharpe ratio is therefore 10 times higher on announcement days; a myopic investor would need to exhibit implausibly high risk aversion in order to account for this difference.

One explanation consistent with our results requires a positive dependence of stock market returns on state variables such as expected long-run economic growth and inflation. Intuitively, stocks tend to perform particularly poorly when news about the state of the economy is negative, making them much riskier than just their volatility would suggest. Given that scheduled economic announcements presumably reveal important information about the economy, this state variable risk should be deterministically higher at such times.[2] Risk premia can therefore

[1]By contrast, we find that the average excess return on days of unscheduled FOMC announcements is strongly negative (−89.2 bp for 15 such days between 2001 and 2009). This is consistent with the prescheduled nature of the announcements in our tests being the crucial characteristic driving differences in risk premia.

[2]In support of this hypothesis, we find that in the 2nd half of our sample period announcement-day returns predict consumption growth significantly better than nonannouncement returns.

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increase on announcement days, even if conditional volatility does not change. This explanation can reconcile the large announcement effect on risk premia with the small effect on observed return volatility and the corresponding difference in Sharpe ratios. Intriguingly, it implies that the major component of the equity premium is compensation for exposure to news about the state of the economy: macroeconomic risk.

This argument also suggests that periods of high uncertainty about the direction of the economy should be times when the difference between announcement- and nonannouncement-day returns is especially high, and we confirm this hypothesis. We estimate that a doubling of stock market variance increases the differential by 60%.

Higher risk on announcement days should also affect the risk-free rate. Increased risk can raise desired saving by risk-averse investors to insure against adverse states of the world. In equilibrium, increased precautionary saving demand should reduce returns on the risk-free asset, and we find strong support for this prediction. The average holding-period return on 30-day U.S. T-bills (our proxy for the daily risk-free rate) is 0.7 bp lower on announcement days, relative to the sample mean of 2.2 bp, with a t-statistic of above 10.

For longer-term Treasury securities, which are not riskless assets at a daily horizon, the difference between average announcement- and nonannouncement-day returns increases monotonically with a bond's maturity, as we would predict if investors expect higher returns on riskier assets on announcement days. Treasury bonds with maturities over 1 year behave similarly to the stock market, with higher excess returns, similar volatilities, and consequently significantly higher Sharpe ratios on announcement days relative to nonannouncement days.

Our results hold over the full 1958–2009 sample (1961–2009 for Treasuries), are almost unchanged in various subsamples, are robust to exclusion of outliers, and hold separately for each type of announcement. They are also not explained by various calendar anomalies, including the January effect, the day-of-the-week effect (French (1980), Gibbons and Hess (1981)), the turn-of-the-month effect (Ariel (1987), Lakonishok and Smidt (1988)), the 1st-half-of-the-month effect (Ariel (1987)), the holiday effect (Ariel (1990)), and seasonality induced by payment lags (Flannery and Protopapadakis (1988)).

It is possible that the postwar period in the United States was an unusually good time for equities, with economic growth consistently beating expectations and a benign inflation experience. If these unexpectedly favorable developments were revealed primarily through scheduled announcements, this could explain our finding that stock returns on announcement days are significantly higher.[3] However, our results remain the same if we include controls for announcement surprises, which we either estimate directly or compute using forecasts by the Society of Professional Forecasters (SPF). Furthermore, the announcement-day premium is present in 9 out of 10 5-year subsamples starting in 1958, some of which do not seem to coincide with positive economic conditions.

[3] An analogous argument can be made for long-term bonds.

A number of papers investigate the *sensitivity* of realized returns to the news component of scheduled macroeconomic announcements. For instance, a positive inflation shock (an announcement of an inflation number higher than the market expectation) may induce a negative contemporaneous stock market return. In the language of factor models, these papers investigate factor betas as opposed to factor risk premia. Formally, given an announcement-day surprise z_{t+1}, defined as the difference between the announced number and its forecast, a test asset return r_{t+1} is decomposed into its conditional expectation and its residual:

$$r_{t+1} = \mathrm{E}_t[r_{t+1}] + \beta z_{t+1} + \varepsilon_{t+1}. \tag{1}$$

Starting with Schwert (1981), Pearce and Roley (1983), (1985), Hardouvelis (1987), Cutler, Poterba, and Summers (1989), Orphanides (1992), McQueen and Roley (1993), Krueger (1996), Fleming and Remolona (1997), Balduzzi, Elton, and Green (2001), Bomfim (2003), and Gurkaynak, Sack, and Swanson (2005) study the responsiveness β of stock or bond returns to various macroeconomic shocks z_{t+1}. More recently, Boyd, Hu, and Jagannathan (BHJ) (2005) explore the sensitivity of security returns to unemployment surprises and find a positive stock market response to news of rising unemployment during economic expansions (a positive β) and a negative response during contractions (a negative β). Andersen, Bollerslev, Diebold, and Vega (2007) use a high-frequency futures data set and get a similar result that the stock market response to macroeconomic news depends on general economic conditions. Bernanke and Kuttner (2005) analyze the impact of FOMC interest rate announcement surprises on stock market returns.

Flannery and Protopapadakis (2002) estimate a direct announcement effect on contemporaneous returns through the sensitivity to announcement news β together with an indirect effect through higher conditional volatility of shocks ε_{t+1} (even if β equals 0) on announcement days. They employ a generalized autoregressive conditional heteroskedasticity (GARCH) model to identify which macroeconomic surprises (out of 17 candidates) influence realized equity returns or their conditional volatility. They come up with 3 variables (CPI, PPI, and the monetary aggregate) for which there exists a relation between surprises and returns, and only one of those (the monetary aggregate) affects returns both directly and indirectly.[4,5]

By contrast, this study focuses on the effect of prescheduled announcements on *expected returns* $\mathrm{E}_t[r_{t+1}]$. We identify the magnitude of the difference between expected returns on announcement days versus expected returns on other days for the stock market, long-term bonds, and T-bills. As a consequence, we are not

[4]The finding that unexpected inflation and money growth negatively affect stock prices is not new. See Bodie (1976), Nelson (1976), Fama and Schwert (1977), Jaffe and Mandelker (1979), Fama (1987), Schwert (1981), Geske and Roll (1983), and Pearce and Roley (1983), (1985) for previous studies establishing this relation.

[5]Brenner, Pasquariello, and Subrahmanyam (2009) estimate a similar GARCH framework for stock, Treasury, and corporate bond markets that allows for an announcement-day effect on the mean through a variance-in-mean channel, but find no evidence of a positive statistically significant effect on average excess returns.

directly interested in the announcement surprise z_{t+1} but rather in the average realized return over a long sample. This means we do not need to make assumptions about market expectations for a given variable or even about what exactly constitutes good or bad news at any particular point in time.[6] We also do not need to know the size or sign of β, as long as we accept the results of the earlier studies that find that β is different from 0, and therefore announcement days are periods of higher systematic risk. Jones, Lamont, and Lumsdaine (1998) adopt a methodology similar to ours and find that both the mean excess returns for long-term Treasury bonds and their volatilities are higher on PPI and employment announcement days.[7]

Our results could be related to the well-known phenomenon of high average stock returns for firms announcing earnings. This earnings-announcement premium was first discovered by Beaver (1968) and was subsequently confirmed by Chari, Jagannathan, and Ofer (1988), Ball and Kothari (1991), Cohen, Dey, Lys, and Sunder (2007), Frazzini and Lamont (2007), and Savor and Wilson (2011), who all find that the above-average returns around earnings announcement days do not appear to be explained by increases in risk. Kalay and Loewenstein (1985) obtain the same finding for firms announcing dividends. While potentially similar, our results are easier to interpret in the framework of a rational choice equilibrium, since we do not need to distinguish between the idiosyncratic component of announcement-day risk and the systematic component. It is not immediately clear to what extent firm-level announcement risk can be diversified away, but macroeconomic announcement risk surely cannot be diversified to any significant extent.

Our explanation for the documented announcement-day premia focuses on a risk-return trade-off that compensates investors for higher macroeconomic risk around announcement days. An interesting alternative possibility is that even myopic investors effectively become more risk averse ahead of announcements, resulting in a higher price of stock market variance. Such investors could be averse to uncertainty in the sense proposed by Knight (1921). With an announcement approaching, their utility functions become more concave as the worst possible distributions of outcomes receive higher weights.[8] Other potential explanations include the changing composition of investors participating in stocks and T-bills ahead of announcements, which would alter the risk aversion of the representative investor, or an irrationally excessive investor aversion to announcement risk.

[6]It is not always obvious how the market will interpret a particular macroeconomic shock. For example, if the stock market response to news of rising unemployment depends on concurrent economic conditions, a lower than anticipated number could represent bad news. Similarly, lower than expected inflation in Japan in recent years was not necessarily good news for investors.

[7]We document a similar result in our sample.

[8]Skiadas (2008) shows that for small risks (a large probability of a small change or a small probability of a large change) many of the preferences in the current literature are, to a 1st-order approximation, equivalent to expected utility or Kreps-Porteous recursive preferences, so that ambiguity aversion need have no 1st-order effects on asset prices when risks are small. Prescheduled announcements, however, are the quintessential large risk: They are events involving the near certainty of a nonnegligible change (even if zero-mean). Thus, even standard ambiguity aversion can deliver higher risk prices ahead of announcements.

The rest of the paper is organized as follows: Section II presents a simple model deriving all of our main predictions; Section III reports our principal results and relates them to our model; Section IV shows additional supporting evidence; and Section V concludes. Derivations of our propositions are given in the Appendix.

II. Announcement Risk in an Endowment Economy

Our intuition is that times around scheduled macroeconomic news announcements are periods of foreseeably higher systematic risk, and that consequently expected excess returns on risky assets should be higher during those periods. In equilibrium, this intuition can also imply that risk-free rates should be lower during the same periods. Here we analyze this idea in a formal model of scheduled announcements in an endowment economy with a single Lucas tree and a single representative investor with recursive preferences, in which inflation and real interest rates are stochastic. All derivations are provided in the Appendix.

The central idea of our model is that investors learn more about the state of the economy on announcement days than on other days. Investors are rewarded not just for bearing market risk but also for state variable risk, which we define as the risk of learning that the economy is performing worse than expected. Consequently, risky assets whose returns have high covariance with the state variable can earn much higher risk premia around announcements, even if the volatility of their returns is not very different. Such assets include the overall stock market and long-term nominal bonds. Since these assets' returns have a larger common component on announcement days, they should comove more around announcements. The model shows how this idea can be made consistent with general equilibrium by equating the state variable in the intertemporal capital asset pricing model of Merton (1973) with long-term expected consumption growth in an endowment economy, as in Bansal and Yaron (2004).

The primary purpose of the model is to demonstrate that all of the announcement effects we document are consistent with a rational expectations equilibrium. The model also makes it easier to understand which forces may be driving our findings and suggests additional testable hypotheses. Most of our empirical work is not directly testing this model; to the extent that it is possible, we do so in Section IV.C.

A. Real Economy

Log real aggregate dividends (which equal the endowment), $d_t = \ln D_t$, follow

(2) $$\Delta d_{t+1} = \mu_t + \nu_{d,t+1}.$$

The expected growth of the endowment (the drift), μ_t, varies randomly over time, following an AR(1) process:

(3) $$\mu_{t+1} = (1-\phi)\overline{\mu} + \phi\mu_t + \nu_{\mu,t+1}.$$

The conditional variances of both news terms are higher on announcement days:

$$(4) \qquad \mathrm{Var}_t[\nu_{x,t+1}] \quad = \quad \sigma^2_{x,L} + (\sigma^2_{x,H} - \sigma^2_{x,L})A_{t+1},$$

for $x = d, \mu$, where A_{t+1} is a deterministic indicator variable that equals 1 if there is a prescheduled announcement between dates t and $t+1$, and 0 otherwise, and $\sigma_{x,H} > \sigma_{x,L}$. The exposition is considerably simplified if we assume that news about current and expected future endowment growth are uncorrelated.

This model is essentially that of Bansal and Yaron (2004) with the addition of deterministic changes in variances due to announcement effects, and we use a similar approximation to solve the model in closed form. Note that the announcement effects on variances are assumed, and the model is used to derive the resulting announcement effects on prices and expected returns.

B. Preferences

A representative investor chooses an optimal consumption path and invests in a claim to the aggregate endowment and a risk-free asset. The investor has recursive Epstein-Zin preferences,

$$(5) \qquad U_t \quad = \quad \left((1-\beta)C_t^{1-\frac{1}{\psi}} + \beta \left(\mathrm{E}_t[U_{t+1}^{1-\gamma}] \right)^{\frac{1-\frac{1}{\psi}}{1-\gamma}} \right)^{\frac{1}{1-\frac{1}{\psi}}},$$

where β is the time discount rate, γ is the coefficient of relative risk aversion, and ψ is the elasticity of intertemporal substitution (EIS). When $\gamma = 1/\psi$, these preferences nest the special case of power utility. Market clearing requires $C_t = D_t$.

We use recursive Epstein-Zin utility rather than the simpler power utility, because in our equilibrium model power utility has some empirically unattractive properties (when risk aversion is greater than 1). Specifically, as noted by Bansal and Yaron (2004), increases in aggregate risk induce an increase in desired precautionary saving, which in equilibrium reduces expected returns on all assets (the wealth effect) and reduces desired portfolio weights on riskier assets (the substitution effect). Assuming investors have power utility preferences requires the wealth effect to dominate the substitution effect, implying that valuations of even risky assets should be increasing in aggregate risk (holding cash flows constant). Furthermore, under power utility, changes in expected consumption growth do not affect risk premia. The more general Epstein-Zin framework avoids these unappealing implications.[9]

[9]See Bansal, Khatacharian, and Yaron (2005) for evidence that both higher aggregate uncertainty and lower expected consumption growth decrease risky asset valuations.

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C. Real Risk-Free Rate

In equilibrium, the investor consumes the aggregate endowment D_t each period, and the risk-free asset is in zero net supply. The equilibrium log risk-free rate is then given by

$$\text{(6)} \quad r_{f,t+1} = -\ln\beta + \frac{1}{\psi}\left(\mu_t + \frac{1}{2}\text{Var}_t\left[\Delta d_{t+1}\right]\right) - \gamma\left(1+\frac{1}{\psi}\right)\frac{1}{2}\text{Var}_t[\Delta d_{t+1}] - \left(\gamma - \frac{1}{\psi}\right)\left(1-\frac{1}{\psi}\right)\text{Var}_t\left[\frac{\rho}{1-\rho\phi}\mu_{t+1}\right].$$

The log risk-free rate consists of 4 terms. The 1st term depends on the rate of time preference. The 2nd depends on the log expected growth rate of consumption. This term is independent of risk aversion γ, but not of risk $\text{Var}_t[\Delta d_{t+1}]$ because of Jensen's inequality: For risk-neutral investors, an increase in the variance of log dividend growth increases the log risk-free rate because log expected dividend growth increases, reducing desired saving.

The 3rd term is a precautionary saving term that is 0 for risk-neutral investors. For risk-averse investors, an increase in aggregate risk raises desired precautionary saving, reducing the market-clearing risk-free rate. The precautionary saving effect of increased risk dominates the effect through the 2nd term if and only if investors are sufficiently willing to substitute consumption across time (increasing in ψ) relative to their willingness to substitute across states (decreasing in γ). A necessary and sufficient condition for the risk-free rate to be decreasing in aggregate risk is that $\psi \geq ((1/\gamma)-1)$. Since ψ is weakly positive, this condition is always fulfilled for investors with greater-than-unit risk aversion.

The 4th term is an additional precautionary saving term proportional to the variance of the permanent component of shocks to expected endowment growth. This term is 0 for both investors with unit elasticities of intertemporal substitution and for investors with power utility. For the case of γ and ψ greater than 1, this term reduces the risk-free rate on announcement days. Risk-averse investors who are highly willing to substitute future for current consumption (those with high ψ) are most prone to changing their desired consumption plans in response to permanent changes in consumption growth. Such investors will wish to save more as the variance of such news increases (holding the risk-free rate constant).

The total precautionary savings effect on the risk-free rate is the sum of the 3rd and 4th terms, and will be higher on announcement days provided γ and ψ are both greater than 1. Since investors are long lived, any such additional desired saving cannot be very large, but even long-lived investors put some weight on smoothing consumption from day to day.

D. Stock Market Returns

The log return on the risky claim to the aggregate endowment is

$$\text{(7)} \quad r_{\text{MKT},t+1} = -\ln\beta + (\gamma-1)\left(1-\frac{1}{\psi}\right)\frac{1}{2}\text{Var}_t\left[\nu_{d,t+1} + \frac{\rho}{1-\rho\phi}\nu_{\mu,t+1}\right] + \frac{1}{\psi}\mu_t + \nu_{d,t+1} + \left(1-\frac{1}{\psi}\right)\frac{\rho}{1-\rho\phi}\nu_{\mu,t+1}.$$

Expected market returns are higher on announcement days provided $(\gamma - 1)(1 - (1/\psi)) > 0$. For the leading empirical case of $\gamma > 1$, this condition requires the EIS ψ to be greater than 1.[10] Investors desire both to increase saving (a wealth effect) when aggregate risk increases and to substitute out of risky into risk-free assets (substitution effect). Provided that $\psi > 1$, the substitution effect will dominate the wealth effect, so that in equilibrium expected returns on risky assets will increase, while those on very low-risk assets will decrease.

The conditional market risk premium is

$$\begin{aligned}
(8) \qquad & \ln \mathrm{E}_t\left[\frac{1+R_{\mathrm{MKT},t+1}}{1+R_{f,t+1}}\right] \\
&= \mathrm{E}_t[r_{\mathrm{MKT},t+1}] - r_{f,t+1} + \frac{\mathrm{Var}_t[r_{\mathrm{MKT},t+1}]}{2} \\
&= -\mathrm{Cov}_t[m_{t+1}, r_{\mathrm{MKT},t+1}] \\
&= \gamma \mathrm{Var}_t[\Delta d_{t+1}] + \left(\gamma - \frac{1}{\psi}\right)\left(1 - \frac{1}{\psi}\right)\mathrm{Var}_t\left[\frac{\rho}{1-\rho\phi}\mu_{t+1}\right] \\
&= \gamma \mathrm{Var}_t[r_{\mathrm{MKT},t+1}] + \frac{\gamma - 1}{\psi}\mathrm{Cov}_t\left[r_{\mathrm{MKT},t+1}, \frac{\rho}{1-\rho\phi}\mu_{t+1}\right].
\end{aligned}$$

It will be higher on announcement days provided ψ is not too low. For the special cases of power utility or unit intertemporal elasticity of substitution, the variance of the permanent component of shocks to economic growth (the 2nd term) does not affect consumption. When both γ and ψ are greater than 1, the market risk premium is increasing in the variance of this permanent component. Thus, market risk premia can be considerably higher on announcement days if investors expect to receive more news about future economic growth on such days.

In this model, the market risk premium is not necessarily proportional to its conditional return variance. Since the stock market return exhibits a positive covariance with permanent shocks to expected economic growth, conservative investors (those with $\gamma > 1$) will demand higher risk premia on announcement days even if the increase in market variance is small. Such investors require compensation for the tendency of the market to perform poorly when news about future economic growth is bad. In consequence, Sharpe ratios can be much higher on announcement days.

E. Nominal Bonds and Inflation

In order to model announcement effect on bonds, we next introduce inflation shocks. The log dollar price of an N-period nominal discount bond is $p^{\$}_{n,t}$, and its real holding-period return is

$$(9) \qquad r_{n,t+1} = p^{\$}_{n-1,t+1} - p^{\$}_{n,t} - \pi_{t+1},$$

[10]Recent work by Bansal and Yaron (2004), Bansal, Tallarini, and Yaron (2008), Vissing-Jorgenson (2002), and others presents evidence and arguments in favor of $\psi > 1$.

where π is the log rate of inflation. We assume

$$\pi_{t+1} = z_t + \eta_{\pi,t+1} \tag{10}$$

and expected inflation z_t follows,

$$z_{t+1} = (1-\lambda)\overline{\pi} + \lambda z_t + \eta_{z,t+1}. \tag{11}$$

Once again, the conditional variances of realized inflation and expected inflation shocks are higher on announcement days. The structural source of inflation and its relation to real variables is beyond the scope of this paper. However, we assume that shocks to realized or expected inflation are not correlated with shocks to realized endowment growth $\nu_{d,t+1}$. The signs of the correlations between expected inflation and expected real endowment growth and between realized inflation and expected endowment growth are discussed below. Following Campbell and Viceira ((2002), Chap. 3), it is helpful to write out the dependencies of the inflation shocks on each other and on shocks to the drift:

$$\eta_{z,t+1} = \beta_{z\mu} \nu_{\mu,t+1} + \varepsilon_{z,t+1} \tag{12}$$

and

$$\eta_{\pi,t+1} = \beta_{\pi\mu} \nu_{\mu,t+1} + \beta_{\pi z}\varepsilon_{z,t+1} + \varepsilon_{\pi,t+1}. \tag{13}$$

The shocks $\nu_{\mu,t+1}$, $\varepsilon_{z,t+1}$, and $\varepsilon_{\pi,t+1}$ are orthogonal but have higher variances on announcement days. The loadings ($\beta_{z\mu}$, $\beta_{\pi\mu}$, and $\beta_{\pi z}$) are assumed to be the same on all days for simplicity. In order to generate a positive inflation risk premium, we require that $\beta_{z\mu}$ be negative, so that shocks to expected inflation are negatively related to shocks to expected economic growth.

F. Nominal Bond Risk Premia

In real terms, consistent with the rest of this section, risk premia on nominal bonds are given by

$$\begin{aligned}
(14)\quad & \mathrm{E}_t[r_{n,t+1}] - r_{f,t+1} + \frac{1}{2}\mathrm{Var}_t[r_{n,t+1}] \\
&= \mathrm{Cov}_t[-m_{t+1}, p_{t+1}^{\$n-1} - p_t^{\$n} - \pi_{t+1}] \\
&= \left(\gamma - \frac{1}{\psi}\right) \frac{\rho}{1-\rho\phi} \mathrm{Var}_t\left[\mu_{t+1}\right] \left(-\frac{1}{\psi}\frac{1-\phi^{n-1}}{1-\phi} - \frac{1-\lambda^{n-1}}{1-\lambda}\beta_{z\mu} - \beta_{\pi\mu}\right).
\end{aligned}$$

The risk premia depend on 3 terms. The 1st term is the risk premium on an N-period real bond. When γ and ψ are both greater or both less than 1, this implies that risk premia are lower on announcement days by an amount increasing in magnitude with bond maturity. Since the short-term real interest rate depends positively on expected endowment growth, and real long-term bond holding-period returns are negatively correlated with the short-term real rate, long-term real bonds offer desirable hedges against the risk of a decline in expected economic growth. As this risk is higher on announcement days, longer-term real bonds should underperform by more on such days.

The 2nd term depends negatively on the covariance between shocks to expected inflation and shocks to expected real endowment growth. In order to generate a positive inflation risk premium, this covariance must be negative. Although there is evidence that the inflation risk premium may have declined over time, most studies agree that it has always been positive (see, e.g., Buraschi and Jiltsov (2007) and Campbell, Sunderam, and Viceira (2011)).

Finally, the 3rd term depends negatively on the covariance between shocks to realized and expected inflation and expected economic growth. The sign of this covariance is a matter of debate, but it is likely to be small in magnitude and is the same for all maturities.

The risk premium on 2 nominal bonds with maturities n+1 and n is increasing in maturity, and higher on announcement days, provided

$$(15) \qquad -\beta_{z\mu} \quad > \quad \frac{1}{\psi}\left(\frac{\phi}{\lambda}\right)^{n-1}.$$

For sufficiently short-term bonds, the risk premium can decline with maturity and will be lower on announcement days. The model therefore predicts that for short-term bonds, the average excess returns on announcement days can be lower than on nonannouncement days, but should always be higher for longer-term bonds.

III. Evidence on Announcement-Day Returns

A. Prescheduled Macroeconomic Announcements

We obtain dates of prescheduled monthly macroeconomic news announcements from the Bureau of Labor Statistics from 1958 to 2009 and from the Federal Reserve from 1978 to 2009. We have 157 prescheduled CPI announcements from Jan. 1958 to Jan. 1971 and 467 for the PPI from Feb. 1971 to Dec. 2009. We drop the CPI after PPI announcements become available in Feb. 1971, since PPI numbers for a given month are always released a few days earlier, thereby diminishing the news content of CPI numbers.[11] We have 621 employment announcements from Jan. 1958 to Dec. 2009. FOMC interest rate announcements start in Jan. 1978 and end in Dec. 2009. Before Feb. 1994, we assume the FOMC decision became public 1 day after its meeting (as in Kuttner (2001)). We exclude any unscheduled announcements, leaving us with 279 FOMC observations. Of the announcement days in our sample, 51 had more than 1 announcement, while a further 23 were nontrading days. The remaining sample contains 1,450 announcement days versus 11,641 nonannouncement days. Interestingly, only 29 of the prescheduled announcements in our sample occurred on a Monday, representing about 2% of overall announcements. In the 2nd half of our sample, there is only 1 Monday announcement.

Our choice of announcement types is primarily dictated by the availability of data. CPI is the 1st macroeconomic variable for which the Bureau of Labor

[11]Our results are robust to the inclusion of CPI announcements after Jan. 1971.

Statistics issued regular news releases (according to data available on its Web site (http://www.bls.gov/data/)), followed 4 years later by employment. We need a long sample for our analysis to ensure that the average surprise is close to 0, so that announcement-day returns do not reflect a period of particularly good or bad news.[12] Moreover, both employment and inflation clearly constitute important macroeconomic news, as do FOMC announcements.[13]

Our measure of stock market return is the daily return on the Center for Research in Security Prices (CRSP) value-weighted NYSE/NASDAQ/AMEX all share index, including dividends. To calculate excess returns, we infer a daily risk-free rate from the monthly risk-free rate (obtained from Kenneth French's Web site (http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/)), assuming it to be constant over the month. This biases downward our estimate of the difference in average excess returns between announcement and nonannouncement days, since we also find evidence consistent with a lower daily risk-free rate on announcement days.

We obtain daily T-bill returns from the CRSP daily Treasuries file starting in June 1961 (the 1st date available) and ending in Dec. 2009. Our proxy for the overnight risk-free rate is the daily return on the T-bill in the CRSP file with maturity closest to 30 days.[14] Our results do not depend on the exact choice of the number of days until maturity.

For Treasury securities with longer maturities, we use returns provided by CRSP's Daily Treasury Fixed Term Indexes File. These returns are meant to reflect the performance of a hypothetical Treasury bond with fixed maturity and are calculated using a procedure similar to the one we employ for calculating our daily risk-free rate.

We obtain constant-maturity 30-day implied volatility from the Chicago Board Options Exchange (CBOE) Standard & Poor's (S&P) 100 Volatility Index (VIX), available daily beginning in 1986. These volatilities are then squared to convert them into variances, and the daily difference from market close to market close is calculated. Estimates of the change in stock market risk based on prices at a point in time such as implied volatilities could be more accurate than estimates based on realized volatility.

B. Stock Market Excess Returns

Table 1 presents our main result: The average excess return on the stock market is 11.4 bp on announcement days versus 1.1 bp on other days. The difference between the returns on the two kinds of days averages 10.3 bp, and a t-test for a difference in means (allowing for different variances) gives a t-statistic of 3.77. The nonannouncement-day returns are not only much lower but are not even statistically significant (t-statistic = 1.29). Excluding outliers (observations outside the 1st and 99th percentiles of each sample), the average excess returns are

[12]A long sample should also address potential critiques based on the peso problem hypothesis.

[13]See Jones et al. (1998), Bernanke and Kuttner (2005), and BHJ (2005) for further evidence of the variables' relevance.

[14]The CRSP file contains very few observations for bonds with initial maturities of less than 6 months. As a result, hardly any of the bills in our sample are on-the-run 30-day T-bills.

11.7 bp and 1.3 bp, respectively, with a t-statistic for different means of 4.55, and the nonannouncement-day returns are only marginally significant (t-statistic = 1.88). This evidence suggests that macroeconomic risks represent important priced factors for stock returns, as the observed equity risk premium is much higher on announcement days.

TABLE 1

Summary Statistics for Daily Stock Market Excess Returns

Table 1 presents the distribution of stock market excess returns on announcement days and nonannouncement days. Announcement days are those trading days when CPI/PPI (CPI before 1971 and PPI afterward) numbers, employment numbers, and FOMC interest rate decisions are scheduled for release. The sample covers the period 1958–2009. Market excess returns are computed as the difference between the CRSP value-weighted market return and the risk-free rate. The daily risk-free rate is derived from the 1-month risk-free rate provided by CRSP. t-statistics are given in square brackets. All numbers are expressed in basis points, and the numbers in bold are of special interest.

	Panel A. All Obs.			*Panel B. Excl. Outliers (<1 percentile or >99 percentile)*		
	Ann.	Nonann.	Diff.	Ann.	Nonann.	Diff.
Mean	11.4	1.1	**10.3**	11.7	1.3	**10.4**
	[4.41]	[1.29]	**[3.77]**	[5.39]	[1.88]	**[4.55]**
1st percentile	−258.2	−257.7	−1.2	−208.4	−203.2	−5.1
25th percentile	−34.7	−40.8	6.1	−33.3	−39.3	6.0
Median	13.5	4.3	9.2	13.5	4.3	9.2
75th percentile	59.0	45.2	13.7	57.8	44.3	13.5
99th percentile	292.6	242.4	50.3	233.4	191.3	42.1
Std. dev.	98.6	94.6		81.8	75.4	
Skewness	−0.6	−0.6		0.0	−0.2	
Kurtosis	8.5	19.7		1.1	0.8	
N	1,450	11,641		1,420	11,407	

Our hypothesis is that announcement days are fundamentally riskier than other days. The standard deviation of announcement-day returns is 98.6 bp versus 94.6 bp for other days (81.8 vs. 75.4 when excluding outliers), and we can reject the hypothesis of equal variances at the 5% significance level. However, the dispersion of announcement-day returns is only 4%–6% higher, so that the Sharpe ratio is about 10 times higher compared to non-announcement days. Furthermore, announcement-day returns exhibit skewness equal to those on other days, and the distribution of announcement-day returns has a thinner left tail than the nonannouncement-day distribution.[15] It appears that announcement days are not riskier simply because the distribution of announcement-day returns is less attractive to a myopic investor. Consequently, if announcement-day risk premia are higher because of higher fundamental risk, this must be because of higher exposure to state variable risk on announcement days.[16]

[15]This remains true even if we exclude the Oct. 1987 market crash, although there is obviously no good reason to exclude such events when evaluating tail risk.

[16]Pollet and Wilson (2010) show that when the stock market is a poor proxy for the portfolio of aggregate wealth, changes in the average correlation between stock returns can nevertheless reveal changes in aggregate risk. We use estimates of daily average correlation based on intraday returns starting in 1995 and find that the mean announcement-day correlation equals 0.245 versus 0.216 on other days (with a t-statistic for the difference of 3.78). This result suggests, as do our findings on realized and implied volatility, that aggregate risk is higher on announcement days, but that the increase is not of the same order of magnitude as the increase in risk premia.

Table 2 presents evidence from regressions of returns on an announcement-day dummy variable together with controls. The regression coefficients are estimated using ordinary least squares (OLS), and t-statistics are computed using Newey-West (1987) standard errors (with 5 lags, but our results do not change with different specifications).[17] Panel A is for the full sample of 13,091 days, and Panel B excludes outliers using the same cutoffs as above. Column 1 of each panel reproduces the difference-in-means result of Table 1: The announcement-day dummy variable has a significantly positive coefficient. We then control for market return lagged 1 day and squared lagged market return. The coefficient on the lagged market return is positive and significant, in accordance with previous work. Finally, we include day-of-the-week dummy variables for Monday–Thursday. The presence of these dummy variables should absorb any impact on returns by different days of the week, which may stem from payment lags, higher or lower trading activity on particular days, or behavioral biases. We confirm that returns are significantly lower on Mondays (even excluding outliers) and otherwise find no significant day-of-the-week effects. The announcement-day effect remains positive and highly significant in all specifications, although slightly lower once day-of-the-week effects are included.

TABLE 2

Regression Analysis: Daily Stock Market Excess Returns

Table 2 presents the results of OLS regressions of daily stock market excess returns on an announcement-day dummy variable and various other controls. Ann. day is a dummy variable equaling 1 if day t is an announcement day, and 0 otherwise. Market excess returns (MKTRF) are computed as the difference between the CRSP value-weighted market return and the risk-free rate (expressed in basis points). Monday–Thursday are dummy variables for the corresponding days of the week. t-statistics are calculated using Newey-West (1987) standard errors (with 5 lags) and are given in square brackets. The numbers in bold are of special interest.

	Panel A. All Obs.			*Panel B. Excl. Outliers (<1 percentile or >99 percentile)*		
Variable	1	2	3	1	2	3
Intercept	1.133	−0.064	3.448	1.324	0.647	4.670
	[1.24]	[−0.07]	[2.00]	[1.71]	[0.83]	[3.11]
Ann. day	**10.291**	**10.439**	**7.093**	**10.38**	**10.653**	**8.462**
	[3.55]	**[3.61]**	**[2.69]**	**[4.18]**	**[4.34]**	**[3.41]**
$MKTRF_{t-1}$		0.081	0.081		0.094	0.094
		[4.91]	[4.93]		[8.75]	[8.74]
$(MKTRF_{t-1})^2$		0.0001	0.0001		0.0000	0.0000
		[1.61]	[1.59]		[0.47]	[0.45]
Monday			−13.605			−11.228
			[−5.05]			[−5.31]
Tuesday			−2.611			−5.662
			[−1.06]			[−2.71]
Wednesday			2.151			1.129
			[0.89]			[0.56]
Thursday			−2.558			−3.519
			[−1.07]			[−1.71]
N	13,091	13,090	13,090	12,827	12,826	12,826
R^2 (%)	0.1	0.9	1.2	0.2	1.4	1.7

[17]Our findings remain unaltered if we instead jointly estimate announcement-day effects on both the mean and conditional volatility (using a GARCH(1,1) model similar to the one used in Jones et al. (1998)). These results are available from the authors.

We do not include the contemporaneous announcement surprise in our controls. We take this approach for two reasons. First, because the correlation between the true announcement surprise and the announcement-day dummy variable (which is deterministic) is 0 by definition, its exclusion does not bias our estimates of the announcement-day effect (whereas the inclusion of a poorly estimated surprise may bias them). Second, the magnitude and even the sign of the response coefficient to the surprise may well vary over time, which complicates any analysis relying on these coefficients. For example, BHJ (2005) show that unemployment shocks have opposite impacts on stock returns in expansions and recessions. In our robustness tests, we show that controls for announcement surprises have no effect on our findings.

C. Risk-Free Rate

Table 3 presents findings on the distributions of announcement- and nonannouncement-day returns on 30-day T-bills. Our sample starts slightly later (1961, rather than 1958), but is otherwise identical to the stock market sample of announcements.

TABLE 3

Summary Statistics for Daily 30-Day T-Bill Returns

Table 3 presents the distribution of daily 30-day T-bill returns on announcement days and nonannouncement days. Announcement days are those trading days when CPI/PPI (CPI before 1971 and PPI afterward) numbers, employment numbers, and FOMC interest rate decisions are scheduled for release. The sample covers the period 1961–2009. The 30-day T-bill returns are defined as the return of the T-bill issue whose length of maturity is closest to 30 days (daily T-bill quotes are obtained from CRSP starting in June 1961). *t*-statistics are given in square brackets. All numbers are expressed in basis points, and the numbers in bold are of special interest.

	Panel A. All Obs.			*Panel B. Excl. Outliers (<1 percentile or >99 percentile)*		
	Ann.	Nonann.	Diff.	Ann.	Nonann.	Diff.
Mean	1.5	2.3	**−0.7**	1.5	2.2	**−0.7**
	[34.69]	[94.84]	**[−14.49]**	[51.29]	[123.03]	**[−21.40]**
1st percentile	−0.9	−0.7	−0.2	−0.3	−0.2	−0.2
25th percentile	0.8	1.0	−0.2	0.8	1.0	−0.2
Median	1.3	1.6	−0.3	1.3	1.6	−0.3
75th percentile	1.9	2.9	−1.0	1.9	2.9	−1.0
99th percentile	6.8	10.2	−3.5	5.5	8.5	−3.0
Std. dev.	1.6	2.5		1.1	1.8	
Skewness	8.5	0.2		1.5	1.5	
Kurtosis	163.1	55.3		3.8	2.3	
N	1,370	10,711		1,342	10,495	

Panel A of Table 3 indicates that the average announcement-day return for 30-day T-bills is 1.5 bp versus 2.3 bp for nonannouncement days. The difference of 0.7 bp is statistically significant with a *t*-statistic of 14.49, and is also economically significant relative to the sample mean of 2.2 bp. The respective standard deviations are 1.6 bp and 2.5 bp. The 30-day T-bill returns are actually less volatile on announcement days, but the main point is that both of these volatilities are extremely small, as one would expect if these returns represented good proxies for the risk-free rate. The distribution of announcement-day returns on 30-day T-bills is everywhere below that of nonannouncement-day returns.

The statistical significance of the result that 30-day T-bill returns are lower on announcement days is stronger if outliers are excluded, with the *t*-statistic for the difference increasing to 21.4. The exclusion of outliers is more important in this case because of the greater possibility of data error, resulting from the fact that the prices of bond trades are not reported to an exchange.

Table 4 gives our regression results. As before, column 1 of Panel A reproduces the difference-in-means result. Column 2 controls for lagged return and lagged squared return. Not surprisingly, T-bill returns are highly autocorrelated, but the announcement-day effect is still highly significant. Column 3 controls for day-of-the-week effects. Returns on T-bills appear to depend strongly on the day of the week, but, even with the inclusion of dummy variables for different days, the announcement-day effect is still very significant (although substantially smaller). We conclude that the evidence is consistent with increased announcement-day risk reducing the risk-free rate.

TABLE 4

Regression Analysis: Daily 30-Day T-Bill Returns

Table 4 presents the results of OLS regressions of daily 30-day T-bill returns on an announcement-day dummy variable and various other controls. Ann. day is a dummy variable equaling 1 if day *t* is an announcement day, and 0 otherwise. The 30-day T-bill returns (TBILL) are defined as the return of the T-bill issue whose length of maturity is closest to 30 days (expressed in basis points). Monday–Thursday are dummy variables for the corresponding days of the week. *t*-statistics are calculated using Newey-West (1987) standard errors (with 5 lags) and are given in square brackets. The numbers in bold are of special interest.

	Panel A. All Obs.			*Panel B. Excl. Outliers (<1 percentile or >99 percentile)*		
Variable	1	2	3	1	2	3
Intercept	2.255	1.936	1.057	2.200	1.902	1.092
	[69.77]	[37.28]	[13.71]	[84.54]	[55.12]	[24.45]
Ann. day	**−0.726**	**−0.707**	**−0.106**	**−0.724**	**−0.694**	**−0.141**
	[−15.42]	**[−15.08]**	**[−2.46]**	**[−22.70]**	**[−22.44]**	**[−4.93]**
$TBILL_{t-1}$		0.131	0.214		0.119	0.209
		[6.35]	[6.37]		[10.6]	[9.04]
$(TBILL_{t-1})^2$		−0.0053	−0.0032		0.0012	0.0017
		[−0.32]	[−0.15]		[1.62]	[0.97]
Monday			2.846			2.647
			[31.89]			[57.48]
Tuesday			−0.124			−0.214
			[−1.18]			[−2.83]
Wednesday			0.250			0.210
			[5.01]			[7.32]
Thursday			0.311			0.280
			[5.89]			[10.09]
N	12,080	11,941	11,941	11,836	11,710	11,710
R^2 (%)	0.9	2.6	22.5	1.7	4.5	36.0

Our regression shows that T-bill returns are extraordinarily high on Mondays. One potential explanation for this is the existence of a weekend effect for T-bills: 3 days pass between the Friday T-bill price observation and the Monday observation, whereas only 1 day passes between all other consecutive price observations (excluding holidays).

In unreported tests, we raise the gross Monday return for T-bills and Treasury bonds to the power of ⅓ and repeat our analysis. (This adjustment is not necessary in the case of stock market returns, as in their case the random component dominates the deterministic component due to the passing of time.) Since Monday is almost never an announcement day, this procedure distinguishes between an announcement-day effect on daily T-bill and bond returns and a mere weekend effect. Crucially, all our findings continue to hold with this adjustment.

D. Treasury Bond Excess Returns

In contrast to T-bills, government securities with longer maturities represent risky assets at a daily horizon. If held to maturity, long-term Treasury bonds will provide a guaranteed (nominal) rate of return, but in the meantime their daily price changes will not be fully predictable and will reflect factors such as changes in interest rates. The possibility of such changes can result in longer-term bonds displaying greater differences between announcement- and nonannouncement-day returns.[18]

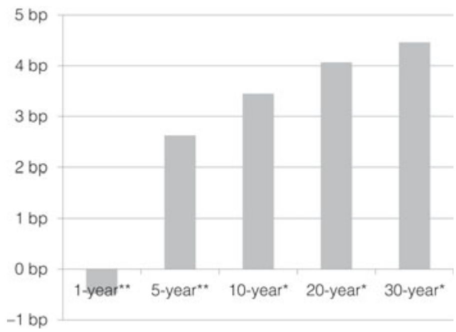
Our model predicts that at long maturities, government bonds should have higher excess returns on announcement days and that the difference should be increasing with maturity, provided that inflation risk premia are positive and shocks to expected inflation are more persistent than shocks to expected economic growth. At the short end of the term structure, it is possible for real interest rate risk premia to dominate inflation risk premia, and thus short-term bond average excess returns can be lower on announcement days (see equation (15)).

This hypothesis is confirmed by the data. Figure 1 shows how the difference between announcement- and nonannouncement-day excess returns varies with a bond's maturity. As predicted, the performance differential uniformly increases as we increase a bond's time to maturity. For a 1-year bond, the average

FIGURE 1

The Difference between Announcement- and Nonannouncement-Day Treasury Bond Excess Returns

Figure 1 plots the difference between the mean announcement-day excess return and the mean excess return on other days for Treasury bonds of different maturities. Treasury bond returns are obtained from the CRSP Fixed Term Indices File. The difference is expressed in basis points (bp). * and ** indicate statistical significance at the 5% and 1% levels, respectively.



[18]For example, simple up or down shifts in the yield curve will have the greatest impact on the Treasury bonds with the longest maturities.

announcement-day excess return is actually 0.5 bp lower than the average on other days, with a t-statistic of 2.22. This suggests 1-year bonds are relatively riskless assets (at a daily horizon). However, as we increase a bond's maturity, its announcement-day returns become higher than nonannouncement-day returns. For 5-year bonds, the return differential is 2.6 bp (t-statistic = 2.57), and it then grows to 3.4 bp (t-statistic = 2.23), 4.1 bp (t-statistic = 2.04), and 4.5 bp (t-statistic = 2.02) for 10-, 20-, and 30-year bonds, respectively. These findings for longer-dated Treasury securities are similar to those reported in Jones et al. (1998) for the period 1979–1995, and are consistent with the hypothesis that investors expect higher returns on riskier assets on days when macroeconomic news is scheduled to be released.

E. Can the Same Parameters Account for All Announcement Effects?

Our model proposes a unified explanation for announcement effects on the risk-free rate, the equity risk premium, and bond risk premia. Having presented our main results, it is now appropriate to ask whether the model provides a quantitatively as well as qualitatively satisfactory explanation. To address this issue, we show and discuss the results of a simple calibration exercise.

Given that equation (6) requires γ and ψ to both be greater than 1 for the risk-free rate effect to exist, we choose other parameters for the model that can match our estimates for the stock market. Specifically, for $\gamma = 1.2$ and $\psi = 1.001$, we can choose other parameters to match the announcement- and nonannouncement-day market average excess returns and volatilities of returns perfectly.[19] For example, setting $\phi = 0.836$, $\sigma_d^A = 15.4\%$, $\sigma_d^N = 15.0\%$, $\sigma_\mu^A = 4.4\%$, and $\sigma_\mu^N = 8 \times 10^{-11}\%$, all quoted on an annualized basis and assuming 250 trading days per annum, generates the required moments.

Are these numbers also consistent with observed bond and T-bill returns? For bond risk premia, we find that we can generate all the observed announcement- and nonannouncement-day average excess returns almost perfectly (with a root mean squared error of 0.25 bp) if we assume $\lambda = 0.833$ (annualized), $\beta_{z\mu} = -1.07$, and $\beta_{\pi\mu} = 0$. Actual (implied) bond average excess returns (expressed in bp) are then: 0.0 (0.1) for 1-year notes; 3.1 (2.8) for 5-year notes; 3.6 (4.0) for 10-year bonds; 4.3 (4.3) for 20-year bonds; and 4.5 (4.5) for 30-year bonds. For nonannouncement-day average excess returns, all are predicted by the model to be 0, which is quite close to the actual estimates of 0.4, 0.5, 0.2, 0.4, and 0.1, respectively. We conclude that the same parameters of the model can match both stock market and bond risk premia (given stock market variance) on both announcement and nonannouncement days for bond maturities of 1 year and above.

With these parameters, the implied difference between the average risk-free rate on announcement days and nonannouncement days is -10.2 bp, as opposed to an actual reduction of -0.7 bp. The model-implied reduction is thus

[19]The value of γ cannot be too high, as a lower bound for σ_μ^N is 0. At that point, equation (8) sets γ equal to the price of market risk on nonannouncement days, which we estimate to be very low. Given a low γ and very similar market variances on both types of day, equation (7) then requires ψ to be close to 1.

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nearly 15 times too large in magnitude. Inspection of equation (6) shows that the announcement-day effect on the risk-free rate contains a term that is equal and opposite to the effect on the stock market risk premium, and that the other terms are likely to be small, given the small observed difference in market return volatility. Since the risk premium increases by over 10 bp on announcement days, the model-implied effect on the risk-free rate is counterfactually large at around −10 bp.

Why might the actual effect on the risk-free rate be far smaller than the one implied by our model? One possible explanation is that ours is a representative agent endowment model in which bonds and bills are assumed to be in zero net supply. When the agent wishes to reduce his holdings of risky assets, there is no other agent to accommodate him, so the risk-free rate must fall to offset his increased demand. In reality, cash, bills and other close substitutes may be much more elastically supplied (relative to equities and long-term bonds), thereby reducing the impact of announcement risk on the risk-free rate relative to the impact on equities or bonds.

F. Announcement Surprises

A possible alternative explanation for our central results is that announcement returns have a high realized Sharpe ratio in our sample because 1958–2009 has been a period of unusually benign economic conditions for the United States, and investors mainly learned this on announcement days. We now evaluate this idea. We note, however, that this explanation implicitly concedes that announcement days are riskier, since these are the days when investors learn the state of the economy. Furthermore, the announcement effect on market returns is present in all but one consecutive 5-year subperiod in our sample, so that investors would have had to be surprised on the upside not just over the entire 1958–2009 period but also in each 5-year subperiod except one.

A straightforward test of this "good-news hypothesis" is to run the following regression for our test assets:

$$R_t = \beta_0 + \beta S_t + \theta A_t + \text{CONTROLS}_t + \varepsilon_t, \tag{16}$$

where R is the asset return, S is the unexpected component of an economic announcement, A is an announcement-day dummy variable, and CONTROLS is a set of controls. If the sample average S is positive, then omitting S might bias upward our estimate of the announcement effect θ.

Any estimate of the announcement surprises $\widehat{S}_t$ from a regression with a nonzero intercept will have a zero in-sample mean. Therefore, we must use pseudo out-of-sample forecast errors for announced inflation and unemployment when estimating regressions. For interest rate announcements, we use federal fund futures, as in Bernanke and Kuttner (2005), and the surprise is then just the difference between the FOMC decision and the futures price-implied interest rate prediction.

To forecast unemployment, we use the model in BHJ (2005), equation (1):

$$
\begin{aligned}
\text{(17)} \qquad \text{UNEM}_t - \text{UNEM}_{t-1} \;=\; & b_0 + b_1\text{IP}_{t-1} + b_2\text{IP}_{t-2} + b_3\text{IP}_{t-4} \\
& + b_4(\text{UNEM}_{t-1} - \text{UNEM}_{t-2}) \\
& + b_5\text{TB3}_t + b_6\text{BA}_t + \varepsilon_t,
\end{aligned}
$$

where UNEM is the unemployment rate, IP is the growth rate of monthly industrial production, TB3 is the change in the 3-month T-bill rate, and BA is the change in the default yield spread between Baa and Aaa corporate bonds (all monthly). We get UNEM and IP from the Bureau of Labor Statistics, and TB3 and BA from Global Financial Data. We compute forecasts using rolling regressions starting in 1953 (so that the forecast in period t is calculated using coefficients estimated with data from 1953 to period $t-1$).

To forecast inflation, we use an IMA(1,1) (integrated moving average) model, as in Nelson and Schwert (1977) (with monthly data) and Stock and Watson (2007) (with quarterly data). Stock and Watson present evidence that this forecast performs as well as any other over our sample period. The moving average coefficient is estimated using a 10-year rolling window of past observations.

Since BHJ (2005) show that the impact of economic news may vary depending on the business cycle, we also include specifications in which we interact surprises with expansion and recession indicators.[20] As an additional check, we repeat our analysis using median forecasts from the SPF, available from the Federal Reserve Bank of Philadelphia and starting in 1981Q3 (for inflation). These are only reported quarterly, so we use the same SPF forecast for each month in a given quarter.

Table 5 reports the results for the stock market (Panel A) and the risk-free rate (Panel B) as dependent variables. In summary, our estimates of the announcement effect on both expected market returns and risk-free rates are unaffected when we include announcement surprises as controls. The announcement dummy coefficient in the stock market return regression is 8.1 (7.6 with no business cycle indicators) when we use forecasts from equation (17) and 11.3 (9.6 with business cycle indicators) when we use SPF forecasts (with the sample starting in 1982 in the latter case). These coefficients are strongly statistically significant and quite close in magnitudes to the estimates with no announcement surprise controls.

In the T-bill regressions, the coefficients on the announcement dummy variables are −0.15 using both model-based and SPF forecasts. Again, they are all highly significant and very similar to those calculated when announcement surprise controls are omitted. The only surprises with significant impact on T-bill returns are FOMC ones, which is exactly what we should expect (it would be quite worrying if T-bill prices did not respond to federal funds rate changes).

Perhaps surprisingly, no announcement surprises have a statistically significant impact on stock market returns in any of our specifications. At first glance,

[20]For example, if negative unemployment surprises represent good news in recessions and bad news in booms, the average news might have been good in our sample even if the average surprise equaled 0.

TABLE 5

Expectations and Announcement Returns

Table 5 presents the results of OLS regressions of daily stock market (Panel A) and Treasury bond (Panel B) excess returns on an announcement-day dummy variable, measures of announcement surprises, and various other controls (same as in previous tables, but not reported). Ann. day is a dummy variable equaling 1 if day *t* is an announcement day, and 0 otherwise. Announcement-day surprise is the difference between actual and expected numbers, and is always interacted with the announcement-day dummy variable. Expectations for inflation and unemployment are calculated either using a model or obtained directly from the Survey of Professional Forecasters (SPF) (available starting 1981Q3), Federal funds rate expectations are computed using federal funds futures, available starting Oct. 1988. Recession dates come from the National Bureau of Economic Research (NBER). *t*-statistics are calculated using Newey-West (1987) standard errors (with 5 lags) and are given in square brackets. The numbers in bold are of special interest.

	Panel A. Stock Market Excess Returns				*Panel B. T-Bill Returns*			
Variable	Model Expect.		SPF Expect.		Model Expect.		SPF Expect.	
Ann. day	**7.607**	**8.076**	**11.278**	**9.638**	**−0.150**	**−0.156**	**−0.145**	**−0.144**
	[2.60]	**[2.75]**	**[2.51]**	**[2.16]**	**[−3.45]**	**[−3.61]**	**[−2.31]**	**[−2.28]**
Inflation Surprise	0.015		0.007		−0.000		0.000	
	[0.82]		[0.37]		[−0.23]		[0.63]	
Unemployment Surprise	0.088		0.151		0.000		0.003	
	[0.36]		[0.62]		[0.11]		[0.54]	
FED Funds Surprise	−1.472		−1.347		−0.039		−0.039	
	[−0.77]		[−0.74]		[−2.16]		[−2.47]	
Infl. Sur. × Expansion		0.013		0.012		0.000		0.000
		[0.88]		[0.55]		[0.33]		[−0.31]
Unemp. Sur. × Expansion		0.310		−0.167		−0.002		−0.001
		[1.28]		[−0.86]		[−0.56]		[−0.12]
FED Sur. × Expansion		−1.915		−1.869		−0.052		−0.049
		[−1.20]		[−1.19]		[−2.82]		[−2.76]
Infl. Sur. × Recession		0.018		−0.002		−0.000		0.000
		[0.43]		[−0.07]		[−0.42]		[1.09]
Unemp. Sur. × Recession		−0.768		1.005		0.013		0.011
		[−1.03]		[1.61]		[1.15]		[1.67]
FED Sur. × Recession		0.161		0.439		0.008		−0.003
		[0.02]		[0.07]		[0.3]		[−0.16]
N	12,987	12,987	7,008	7,008	11,841	11,841	6,864	6,864
R^2 (%)	1.2	1.2	0.3	0.4	22.6	22.6	14.7	14.7

this seems to conflict with the results in BHJ (2005) (for unemployment) and Bernanke and Kuttner (2005) (for FOMC decisions). However, when we restrict our sample to the same period as in BHJ (1972–2000), we get results that are similar in magnitude and statistical significance. We can also replicate the results in Bernanke and Kuttner when we add unscheduled FOMC announcements to our sample. It seems that their findings are driven exclusively by interest rate changes that occurred outside of the regular FOMC meeting schedule.

In unreported tests, we find that our results for Treasury bonds also hold when we add announcement surprises as controls. Taken together with our finding that stock market announcement effects exist in all but one 5-year subperiod in our sample, we conclude that it is unlikely that our results can be explained by good average announcement news during the period 1958–2009.

G. Subsamples and Other Robustness Tests

Our findings hold separately in each half of the sample, with the difference between announcement- and nonannouncement-day stock market returns (8.7 bp in the 1958–1983 period vs. 11.4 bp in the 1984–2009 period) and risk-free rates

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(−0.7 bp and −0.6 bp) almost the same across the 2 subsamples. This further strengthens the case that the announcement-day premium is not a temporary phenomenon or a chance occurrence. The results are also present for each type of announcement. When we divide the sample further into consecutive 5-year periods, the stock market excess return is higher on announcement days in 9 out of 10 periods, and the T-bill returns are lower in 9 out of 10 periods.

The announcement-day returns are higher for all 10 Fama-French (http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/Data_Library/det_10_ind_port.html) industry portfolios, with the difference being statistically significant for every industry except Durables and Telephone and Television Transmission. Finally, various calendar anomalies, such as the January effect, the turn-of-the-month effect (high equity returns over a 4-day interval beginning with the last trading day of the month, first discovered by Ariel (1987) and Lakonishok and Smidt (1988)), the 1st-half-of-the-month effect (positive stock returns only during the 1st half of calendar months, as in Ariel (1987)), the holiday effect (good stock market performance ahead of market holidays, documented by Ariel (1990)), or seasonality in returns induced by payment lags (Flannery and Protopapadakis (1988)), do not explain any of our results.[21,22]

IV. Additional Tests and Other Supporting Evidence

In this section we show additional results on announcement-day effects.[23] We present evidence that the announcement-day excess return increases with conditional stock market variance, while there is no corresponding effect for nonannouncement days (i.e., at times of high uncertainty, the announcement-day risk premium is higher); that stock market implied variance is higher immediately before announcements; that announcement-day returns predict consumption growth better than nonannouncement-day returns, though only in the 2nd half of our sample; and that the stock market betas of government bonds are much higher on announcement days and the difference in betas is increasing with maturity.

A. Conditional Variance and Stock Market Excess Returns

Our basic argument is that investors demand compensation for the higher risk of learning bad news about the state of the economy on announcement days, resulting in a higher expected return on these days. In times of high uncertainty, the risk of learning bad news is higher than in normal times. Consequently, the differential between returns on announcement and nonannouncement days should be greater during such periods. It is interesting to note that announcement-day returns were on average particularly high during the crisis period in 2008 and 2009,

[21] We control for these effects by introducing dummy variables for each calendar month, a dummy variable for the 4 trading days around the turn of the month, a dummy variable for the 1st half of a month, dummy variables for trading days just before and just after a holiday, and holiday dummy variables interacted with day-of-the-week dummy variables.

[22] All of these results are available from the authors.

[23] All of these findings are implied by our model, but we do not formally derive every prediction.

even though some of the worst announcement-day returns in our sample occurred then.

In more formal terms, the higher Sharpe ratio of announcement-day stock market returns indicates that most of the difference between announcement and nonannouncement returns is due to higher state variable risk and not to higher market variance. It is then simple to show that this difference is increasing in the conditional variance of the market return.[24]

To test this hypothesis, we run the same regression specification as in Table 2 but add as dependent variables the lagged realized variance of stock returns and its interaction with the announcement-day dummy variable. We use the realized variance (over the last 100 trading days) as a proxy for the conditional variance, as in Campbell, Lettau, Malkiel, and Xu (2001). The coefficient on the interaction term (using demeaned variance) is significantly positive with a t-statistic of 2.22. By contrast, the variance coefficient on its own is negative and not statistically significant. The point estimates imply that a doubling of stock market variance from its sample mean increases the announcement-day risk premium by 6.0 bp (relative to the sample mean of 11.4 bp), while not affecting (or even reducing if we use the insignificant coefficient on the realized variance) the nonannouncement-day risk premium.[25] This evidence is consistent with our model and further supports the hypothesis that announcement days are fundamentally riskier.

B. Implied Variance

Our model predicts a drop in VIX, or other Black-Scholes (BS) (1973) implied volatility measures, from before to after announcements. Intuitively, one can think of 30-day-ahead VIX as a "portfolio" of 1-day conditional volatilities. When a high-volatility day, such as an announcement day, drops out and is replaced by a low-volatility one, the "portfolio" volatility drops. We present results on squared implied volatility (implied variance), as these are slightly easier to interpret.

Panel A of Table 6 gives summary statistics for the percentage change in implied variance from previous-day market close to following-day market close, and compares the changes on announcement days to those on nonannouncement days. The average announcement-day change is -1.5%, whereas for other days the average change is an increase of 1.4%. Both estimates are statistically significant, and the difference is large and highly statistically significant (t-statistic $=$ 4.30). The median change in implied variance around nonannouncement days is effectively 0. The median change around announcement days is -2.9%, and the distribution of announcement-day changes lies everywhere below the distribution of nonannouncement-day changes. When we exclude outliers in Panel B, our findings remain the same and become even more significant.

In untabulated results, we also run a regression of implied variance on an announcement dummy variable with controls for lagged changes in implied

[24]See Appendix, equation (A-19).

[25]The mean announcement-day variance is 9586 bp^2. The interaction coefficient is 0.00077 and the variance coefficient is -0.00014. A doubling of variance then increases the expected excess return by $9586 \times 0.00077 - 9586 \times 0.00014$.

TABLE 6

Summary Statistics for Implied Variance

Table 6 presents the distribution of daily changes in implied variance on announcement days and nonannouncement days. Announcement days are those trading days when CPI/PPI (CPI before 1971 and PPI afterward) numbers, employment numbers, and FOMC interest rate decisions are scheduled for release. Implied variance is calculated from the CBOE S&P 100 Volatility Index, which is a constant-maturity 30-day measure of the expected volatility for the S&P 100 Index (available starting in 1986). The daily change is computed as the difference between the end-of-day value on that day and the end-of-day value on the previous trading day. *t*-statistics are given in square brackets. All numbers are expressed in percentage points, and the numbers in bold are of special interest.

	Panel A. All Obs.			*Panel B. Excl. Outliers (<1 percentile or >99 percentile)*		
	Ann.	Nonann.	Diff.	Ann.	Nonann.	Diff.
Mean	−1.5	1.4	**−2.8**	−2.2	0.8	**−3.0**
	[−2.64]	[3.89]	**[−4.30]**	[−5.47]	[5.52]	**[−7.06]**
1st percentile	−28.3	−26.0	−2.3	−25.8	−21.7	−4.1
25th percentile	−9.3	−6.2	−3.1	−9.1	−6.1	−3.1
Median	−2.9	−0.1	−2.8	−2.9	−0.1	−2.8
75th percentile	3.2	6.6	−3.4	3.1	6.4	−3.3
99th percentile	53.1	41.0	12.1	31.9	33.4	−1.5
Std. dev.	15.4	25.5		10.8	10.7	
Skewness	4.1	47.1		0.8	0.6	
Kurtosis	36.7	2,944		1.9	1.0	
N	754	5,289		738	5,183	

variance, the square of such lagged changes, and day-of-the-week dummy variables. None of the results in Table 6 are materially affected. In sum, our evidence strongly suggests that the implied variance falls after macroeconomic news is released. Ederington and Lee (1996) obtain a similar result for interest rate options, while Dubinsky and Johannes (2005) document a decline in implied volatility for individual stock options after earnings announcements.[26]

C. Covariance with Consumption Growth

Our explanation for the much higher Sharpe ratio of announcement-day stock returns is that investors face higher state variable risk on these days. In our model, this state variable risk is equated to the risk of learning that future economic growth, and therefore consumption growth, will be lower than expected. A literal test of our model would then check whether stock market announcement-day returns predict future consumption growth better than nonannouncement-day returns.

Table 7 presents the results of this test. Our measure of consumption is the log real per capita consumption based on National Income and Product Accounts (NIPA) consumption data, taken from Martin Lettau's Web site (http://faculty.haas.berkeley.edu/lettau/data.html) and available at a quarterly frequency. We include 3 lags of consumption growth as controls (we find that only the first 3 are significant). We add all the announcement-day excess returns in a given

[26]Beber and Brandt (2009) use prices of economic derivatives to measure macroeconomic uncertainty and show that implied volatilities of stock and bond options decline more after news releases when uncertainty is high. Li and Engle (1998) and French, Leftwich, and Uhrig (1989) show a reduction in return volatility immediately prior to a prescheduled announcement for the bond and agricultural futures markets, respectively.

quarter t to obtain cumulative announcement-day returns, and we do the same for nonannouncement-day excess returns. Their sum, $r_{q,t}^A + r_{q,t}^N$, is the cumulative market excess return in quarter t.

TABLE 7

Regression Analysis: Consumption Growth and Announcement-Day Excess Returns

Table 7 presents the results of OLS regressions of quarterly log real per capita aggregate consumption growth on lagged cumulative quarterly announcement-day (r_q^A) and nonannouncement-day (r_q^N) excess returns, plus 3 lags of consumption growth. Consumption data come from Martin Lettau's Web site (http://faculty.haas.berkeley.edu/lettau/data.html). Cumulative quarterly returns are computed by summing up returns over a given quarter. t-statistics are calculated using Newey-West (1987) standard errors (with 5 lags) and are given in square brackets. The numbers in bold are of special interest.

Variable	1958–2009	1958–1983	1984–2009	1958–2009	1958–1983	1984–2009
Intercept	0.0019	0.0030	0.0010	0.0019	0.0030	0.0010
	[3.84]	[4.31]	[1.60]	[3.84]	[4.30]	[1.60]
$r_{q,t-1}^A$	0.0184	0.0028	0.0213			
	[2.34]	[0.11]	[2.75]			
$r_{q,t-1}^N$	0.0101	0.0163	0.0053			
	[2.84]	[2.97]	[1.28]			
C_GROWTH$_{t-1}$	0.2478	0.1842	0.3107	0.2478	0.1842	0.3107
	[3.47]	[2.08]	[3.33]	[3.47]	[2.08]	[3.33]
C_GROWTH$_{t-2}$	0.1341	0.1567	0.0581	0.1341	0.1567	0.0581
	[1.76]	[1.66]	[0.56]	[1.76]	[1.66]	[0.56]
C_GROWTH$_{t-3}$	0.1882	0.1242	0.3119	0.1882	0.1242	0.3119
	[3.06]	[1.73]	[3.53]	[3.06]	[1.73]	[3.53]
$r_{q,t-1}^A + r_{q,t-1}^N$				0.0142	0.0096	0.0133
				[3.08]	[0.75]	[2.54]
$r_{q,t-1}^A - r_{q,t-1}^N$				**0.0042**	**−0.0068**	**0.0080**
				[1.05]	**[−0.47]**	**[2.39]**
N	206	103	103	206	103	103
R^2 (%)	25.7	20.8	38.0	25.7	20.8	38.0

The first 3 columns of Table 7 present the results of regressing quarterly consumption growth in quarter t on quarter $t-1$ cumulative announcement and nonannouncement returns for the full sample (consumption data end in 2009Q3), the 1st half of the sample (1958Q1–1983Q4), and the 2nd half of the sample (1984Q1–2009Q3). In the full sample, both types of returns forecast future consumption growth. The coefficient on announcement-day returns is 0.018 (t-statistic $=$ 2.34) versus 0.010 for nonannouncement-day returns (t-statistic $=$ 2.84). In the 1st half, only nonannouncement-day returns are significant (t-statistic $=$ 2.97), while in the 2nd half, only announcement-day returns are significant (t-statistic $=$ 2.75).

The next 3 columns of Table 7 run the same regressions but replace the levels of returns with the difference between cumulative announcement- and nonannouncement-day returns, $r_{q,t}^A - r_{q,t}^N$, and their sum, the cumulative market return. If the coefficient on the difference is positive, we can then reject the null that announcement-day returns predict future consumption growth equally or less well than nonannouncement-day returns. The coefficient on the difference is not significant in either the full sample or in the 1st half. However, in the 2nd half, announcement-day returns predict future consumption growth significantly better than returns on nonannouncement days, with a coefficient on $r_{q,t}^A - r_{q,t}^N$ equaling 0.008 (t-statistic $=$ 2.39). This finding is consistent with investors facing a higher risk of learning bad news about future consumption on announcement days. It is important to emphasize here that in a typical quarter there are 55 nonannouncement days and only 8 announcement days, but cumulative announcement-day returns still offer more information about the next quarter's consumption growth despite this preponderance of nonannouncement days.

Why does our prediction about consumption growth only hold in the 2nd half? An in-depth investigation of this question is beyond the scope of the current paper, but we suggest three possible explanations, all of which are also consistent with the finding that R^2 is almost twice as high in the 2nd half. First, consumption data may be of poorer quality in the earlier period. Wilcox (1992) shows that NIPA data are likely to contain significant sources of error. Second, the stock market participation rate has increased over time and is much higher in the 2nd half than the 1st half (see, e.g., Bertaut and Starr-McCluer (2000) and Hong, Kubik, and Stein (2004)). As argued by Mankiw and Zeldes (1991) and Parker (2001), it is only stockholder consumption that should covary with stock market returns. Third, inflation was more volatile in the early half of our sample, and investors may have overestimated real consumption growth as a result of underestimating inflation in that period. Campbell and Vuolteenaho (2004) argue that inflation illusion is a significant source of potential stock mispricing, and in particular that stock market participants appear to underestimate future real dividend growth when inflation is high, as hypothesized by Modigliani and Cohn (1979).[27]

D. Bond Betas

Table 8 presents betas of government bonds with respect to the stock market return. We regress the excess return of Treasury bonds with different maturities on the stock market excess return, the announcement-day dummy variable, and the interaction term between the two. The coefficient on the announcement-day dummy variable corresponds to the chart in Figure 1: It is negative for the shortest horizon (t-statistic $= -2.71$) and then becomes positive for a 5-year horizon (t-statistic $= 2.41$) and continues increasing monotonically with bond maturity. While 1-year bonds underperform on announcement days, those with longer maturities outperform, and this outperformance increases as maturity goes up.

We observe a similar pattern for bond betas. The interaction term, which measures the difference between bond betas on announcement and nonannouncement days, is always positive and significant, and it increases with the maturity of the bond. The difference is 0.010 (t-statistic $= 4.26$) for 1-year bonds, and it then monotonically rises to 0.112 (t-statistic $= 5.79$) for 30-year bonds. With the exception of those with a 20-year maturity, Treasury bonds do not move together with the stock market on nonannouncement days. In contrast, on announcement days this comovement is always significantly positive.

[27]Consistent with this explanation, we find that in the 1st half of our sample, in a similar test but replacing real with nominal consumption growth, cumulative announcement-day returns predict nominal consumption growth better than nonannouncement-day returns (with a marginally significant t-statistic of 1.63). In the 2nd half, only real consumption growth is better predicted by announcement-day returns. These results are available from the authors.

TABLE 8

Bond Betas on Announcement and Nonannouncement Days

Table 8 presents the results of OLS regressions of daily excess returns for Treasury bonds with different maturities on contemporaneous stock market excess returns and an announcement-day dummy variable. Ann. day is a dummy variable equaling 1 if day t is an announcement day, and 0 otherwise. The sample covers the period 1961–2009. Market excess returns (MKTRF) are computed as the difference between the CRSP value-weighted market return and the risk-free rate (expressed in basis points). Treasury bond returns are obtained from the CRSP Fixed Term Indices File (expressed in basis points). t-statistics are given in square brackets. The numbers in bold are of special interest.

	Bond				
Variable	1-Year	5-Year	10-Year	20-Year	30-Year
Intercept	0.443	0.489	0.208	0.351	0.118
	[6.15]	[1.65]	[0.46]	[0.58]	[0.18]
MKTRF_t	−0.001	−0.002	0.007	0.019	0.007
	[−1.00]	[−0.53]	[1.39]	[2.97]	[1.04]
Ann. day	**−0.583**	**2.132**	**2.492**	**2.731**	**3.054**
	[−2.71]	**[2.41]**	**[1.85]**	**[1.51]**	**[1.57]**
Ann. day × MKTRF_t	**0.010**	**0.042**	**0.075**	**0.097**	**0.112**
	[4.26]	**[4.78]**	**[5.61]**	**[5.38]**	**[5.79]**
N	12,081	12,081	12,081	12,081	12,081
R^2 (%)	0.2	0.3	0.4	0.5	0.4

This evidence is consistent with the existence of a priced common factor to stock and bond returns on announcement days that is less present at other times. It is also predicted by our model if the announcement-day increase in the variance of news about expected future consumption growth is greater than the announcement-day increase in the variance of news about current growth. In other words, provided the information that arrives specifically on announcement days is more relevant to state variables such as expected economic growth or expected inflation, as opposed to realized economic growth or realized inflation, bonds and stocks should comove more around announcements.

This point is perhaps most easily understood by considering an extreme but empirically plausible case. Suppose: i) the only sources of time variation in expected returns are expected economic growth and expected inflation; ii) investors learn nothing about current growth through announcements and nothing about expected future growth or inflation (and, by implication, interest rates) other than through announcements; iii) shocks to expected inflation are negatively correlated with shocks to expected economic growth; and iv) shocks to realized inflation and economic growth are independent of each other and of everything else. Since bond returns depend only on news about nominal interest rates, bond returns will be deterministic on nonannouncement days, and their market betas will be 0. On announcement days both the market return and bond returns will respond negatively to news that future inflation will be higher than anticipated, so bond betas will be positive and increasing with maturity.

V. Conclusion

We show that average excess returns and Sharpe ratios on the U.S. stock market are much higher on days when important macroeconomic news is scheduled to be announced. This difference is especially pronounced at times of high risk. We also find that returns on 30-day T-bills, our measure of the risk-free rate,

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are significantly lower on these days. For longer-term Treasury securities, which are not riskless assets at a daily horizon, we find that the difference between announcement- and nonannouncement-day returns uniformly increases with a bond's maturity and is positive for bonds with maturities of 5 years or more. Bonds comove much more with the stock market on announcement days, and this tendency also monotonically increases with maturity.

Our results demonstrate a clear link between macroeconomic risk and financial asset returns. Investors seem to require higher expected returns on risky assets as a compensation for bearing risks associated with macroeconomic news. In addition, the risk premium on nonannouncement days appears to be very low, with our numbers implying that over 60% of the cumulative annual excess return for the stock market is earned on announcement days.

Our findings on risk-free rates are consistent with precautionary saving. If aggregate risk is higher on announcement days, then investors who care about daily changes in their wealth will seek to save more out of current wealth on those days relative to other days. To our knowledge, this is some of the first evidence of precautionary saving affecting asset prices.

These results are consistent with a simple equilibrium model of economy-wide risk that varies deterministically over time because of prescheduled announcements. This model can reconcile the large increase in stock market risk premia with the relatively small increase in stock market variance that we estimate. Because investors learn more about future economic conditions around announcements, they should be less willing to hold assets, such as stocks, that covary positively with these news, even if the variance of their returns is itself not much higher. If such shocks are persistent, even a small increase in their volatility (the news arrival rate) around announcements can result in large increases in the market risk premium. A reasonable calibration of our model produces risk premia and volatilities that match our empirical results. We also provide direct evidence that investors learn more about future consumption growth on announcement days, though this finding is only significant in the 2nd half of our sample.

Appendix. Proposition Derivations

1. Proof of Equations (6), (7), and (8)

For a representative investor with Epstein-Zin preferences, the stochastic discount factor is given by

$$m_{t+1} \quad = \quad \ln M_{t+1} \quad = \quad \theta \ln \beta - \frac{\theta}{\psi} \Delta d_{t+1} - (1-\theta) r_{\mathrm{MKT},t+1}, \tag{A-1}$$

where r_{MKT} is the log return on the market portfolio, defined as the claim to aggregate dividends in perpetuity, and $\theta = (1-\gamma)/(1-(1/\psi))$.

Since everything is log-normal, the log return on any asset $r_{j,t+1}$ is then given by

$$\mathrm{E}_t[m_{t+1} + r_{j,t+1}] + \frac{1}{2}\mathrm{Var}_t[m_{t+1} + r_{j,t+1}] \quad = \quad 0. \tag{A-2}$$

In order to solve the model, we use the Campbell and Shiller (1988) approximation for the log return on the market portfolio

$$r_{\mathrm{MKT,t+1}} \quad \approx \quad k + \Delta d_{t+1} + \rho(p_{t+1} - d_{t+1}) - (p_t - d_t), \tag{A-3}$$

where k is an unimportant constant and $\rho = (1 + \exp(\overline{d-p}))^{-1}$ is another constant that is slightly less than 1. We assume that announcements are not spaced through our sample in such a way that the mean log dividend-price ratio is badly defined. A sufficient condition is that announcements are regularly spaced, so that in any long period, such as 1 year, there is a fixed number.

Next we assume that the log aggregate price-dividend ratio is linear in the drift term μ_t, and its intercept is a deterministic function of time:

$$p_t - d_t = a_{0,t} + a_1 \mu_t. \tag{A-4}$$

As in Bansal and Yaron (2004), a_1 is positive and the price-dividend ratio is increasing in expected dividend growth if and only if $\psi > 1$, so that the direct effect on wealth through increased growth more than offsets the indirect effect through a higher discount rate due to higher expected growth.

The solution implies that the stochastic discount factor is given by

$$m_{t+1} = -\delta_{t+1} - \frac{1}{\psi}\mu_t - \gamma v_{d,t+1} - \left(\gamma - \frac{1}{\psi}\right)\frac{\rho}{1-\rho\phi} v_{\mu,t+1}, \tag{A-5}$$

where

$$\delta_{t+1} = -\ln\beta - (1-\gamma)\left(\gamma - \frac{1}{\psi}\right)\frac{1}{2}\mathrm{Var}_t\left[\nu_{d,t+1} + \frac{\rho}{1-\rho\phi}v_{\mu,t+1}\right]. \tag{A-6}$$

Iterating equation (A-4) forward 1 period gives

$$p_{t+1} - d_{t+1} = a_{0,t+1} + a_1\mu_{t+1}. \tag{A-7}$$

Plugging these into the approximation (A-3) for the log market portfolio return, then plugging the derived expression into the pricing equation (A-2), given the equation for the stochastic discount factor (A-1), and equating coefficients gives

$$a_1 = \frac{1 - \frac{1}{\psi}}{1-\rho\phi} \tag{A-8}$$

and

$$a_{0,t} = b_0 + b_1 A_{t+1} + \rho a_{0,t+1}, \tag{A-9}$$

confirming our conjecture. Here,

$$\begin{aligned} b_0 = {} & \ln\beta + k + \rho a_1(1-\phi)\overline{\mu} \\ & -\frac{1}{2}(\gamma-1)\left(1-\frac{1}{\psi}\right)\left(\sigma^2_{d,L} + \left(\frac{\rho}{1-\rho\phi}\right)^2\sigma^2_{\mu,L}\right) \end{aligned} \tag{A-10}$$

and

$$b_1 = -\frac{1}{2}(\gamma-1)\left(1-\frac{1}{\psi}\right)\left(\left(\sigma^2_{d,H} - \sigma^2_{d,L}\right) + \left(\frac{\rho}{1-\rho\phi}\right)^2\left(\sigma^2_{\mu,H} - \sigma^2_{\mu,L}\right)\right). \tag{A-11}$$

Assuming no rational bubbles implies

$$\lim_{s\to\infty}\rho^s \mathrm{E}_t[p_{t+s} - d_{t+s}] = 0, \tag{A-12}$$

hence

$$\begin{aligned} \lim_{s\to\infty}\rho^s a_{0,t+s} + \lim_{s\to\infty}\rho^s a_1 \mathrm{E}_t[\mu_{t+s}] &= \lim_{s\to\infty}\rho^s a_{0,t+s} + a_1\overline{\mu}\lim_{s\to\infty}\rho^s \\ &= \lim_{s\to\infty}\rho^s a_{0,t+s} = 0, \end{aligned} \tag{A-13}$$

hence

$$a_{0,t} = \frac{b_0}{1-\rho} + b_1 \sum_{j=1}^{\infty} \rho^j A_{t+1+j}. \tag{A-14}$$

Plugging back into the approximation (A-3) gives equation (7). Equation (6) follows from substituting equation (7) into expression (A-1). Subtracting equation (6) from equation (7) gives equation (8).

2. Derivation of Nominal Bond Risk Premia

The price of a nominal bond is derived by conjecturing that

$$p^{\$}_{n,t} = c^n_{0,t} + c^n_1 \mu_t + c^n_2 z_t, \tag{A-15}$$

where $c^n_{0,t}$ is a deterministic function of time and maturity and the other coefficients depend only on maturity. Since the log price of \$1 is 0, all coefficients equal 0 at $n = 0$. Since the bond's real return is $p^{\$}_{n-1,t+1} - p^{\$}_{n,t} - \pi_{t+1}$, iterating forward, plugging the conjecture into equation (A-2) and equating coefficients confirms the conjecture and in particular gives

$$c^n_1 = -\frac{1}{\psi}\frac{1-\phi^n}{1-\phi} \quad \text{and} \tag{A-16}$$

$$c^n_2 = -\frac{1-\lambda^n}{1-\lambda}. \tag{A-17}$$

Equations (14) and (15) follow.

3. Changes in Stock Market Variance

Equation (8) holds both on announcement days and nonannouncement days. The difference in market risk premium $rp_{\mathrm{MKT,t}} = \ln \mathrm{E}_t\left[(1+R_{\mathrm{MKT,t+1}})/(1+R_{f,t+1})\right]$ is therefore

$$\begin{aligned} rp^A_{\mathrm{MKT},t} - rp^N_{\mathrm{MKT},t} &= \gamma\left(\mathrm{Var}_t[r^A_{\mathrm{MKT,t+1}}] - \mathrm{Var}_t[r^N_{\mathrm{MKT},t+1}]\right) \\ &\quad + \frac{\gamma-1}{\psi}\left(\mathrm{Cov}_t\left[r^A_{\mathrm{MKT},t+1}, \frac{\rho}{1-\rho\phi}\mu^A_{t+1}\right] - \mathrm{Cov}_t\left[r^N_{\mathrm{MKT},t+1}, \frac{\rho}{1-\rho\phi}\mu^N_{t+1}\right]\right) \\ &\approx \frac{\gamma-1}{\psi}\left(\mathrm{Cov}_t\left[r^A_{\mathrm{MKT},t+1}, \frac{\rho}{1-\rho\phi}\mu^A_{t+1}\right] - \mathrm{Cov}_t\left[r^N_{\mathrm{MKT},t+1}, \frac{\rho}{1-\rho\phi}\mu^N_{t+1}\right]\right). \end{aligned} \tag{A-18}$$

The approximation follows because the difference in variance of stock market returns between announcement and nonannouncement days is negligible, so that

$$\sigma^A_{\mathrm{MKT},t} = \sqrt{\mathrm{Var}_t[r^A_{\mathrm{MKT},t+1}]} \approx \sigma^N_{\mathrm{MKT},t} = \sigma_{\mathrm{MKT},t}.$$

Now assume (for simplicity) that the correlation $\rho_{\mathrm{MKT},\mu}$ is constant and does not vary across days. Defining $\sigma^2_{\mu,t} = \mathrm{Var}_t[\mu_{t+1}]$,

$$rp^A_{\mathrm{MKT},t} - rp^N_{\mathrm{MKT},t} = \frac{\gamma-1}{\psi}\rho_{\mathrm{MKT},\mu}\sigma_{\mathrm{MKT},t}\left(\sigma^A_{\mu,t} - \sigma^N_{\mu,t}\right), \tag{A-19}$$

which is increasing in $\sigma^2_{\mathrm{MKT},t}$.

4. Stock Market Implied Volatility

Our model has implications for BS (1973) implied volatilities, such as the CBOE's (old) VIX. Under the assumptions of the BS model, the square of the implied volatility of a τ-day option (assuming no dividends are paid between dates t and $t+\tau$) is the conditional variance of the log τ-day-ahead price $p_{t+\tau}$:

$$\text{(A-20)} \qquad \sigma^2_{\tau,\mathrm{BS}} = \mathrm{Var}_t[\ln P_{t+\tau}] = \mathrm{Var}_t\left[p_{t+\tau}\right].$$

Since

$$\text{(A-21)} \qquad p_{t+\tau} = \left(p_{t+\tau} - d_{t+\tau}\right) + d_{t+\tau},$$

the BS implied variance is approximately

$$\text{(A-22)} \qquad \sigma^{\tau 2}_{\mathrm{BS},t} \approx \left(\frac{1-\frac{1}{\psi}}{1-\rho\phi}\right)^2 \mathrm{Var}_t\left[\sum_{j=1}^{\tau}\phi^{\tau-j}\nu_{\mu,t+j}\right] + \mathrm{Var}_t\left[\sum_{j=1}^{\tau}\nu_{d,t+j}\right].$$

The model-implied change in the square of constant-maturity BS (1973) implied volatility from the day prior to an announcement to the end of the following day is therefore

$$\text{(A-23)} \qquad \Delta\sigma^{\tau 2}_{\mathrm{BS},t+1} = \left(\sigma^2_{d,H} - \sigma^2_{d,L}\right)\left(A_{t+1+\tau} - A_{t+1}\right) + \left(\sigma^2_{\mu,H} - \sigma^2_{\mu,L}\right)\left(\frac{1-\frac{1}{\psi}}{1-\rho\phi}\right)^2 \sum_{j=1}^{\tau}\left(\phi^{\tau-j}\right)^2\left(A_{t+1+j} - A_{t+j}\right),$$

where τ is the number of days until expiration of the options from whose prices the implied volatility is derived. In the case of VIX, τ is standardized to 30 days and is quoted on an annualized basis, so it will change by $(365/30)\Delta\sigma^{\tau 2}_{\mathrm{BS},t+1}$ from date t to date $t+1$.

This change consists of 2 terms. First, if date $t+1$ is an announcement day and date $t+31$ is not, then squared implied volatility will decline by an amount equal to the increase in variance of dividend growth around announcement days. Intuitively, one can think of VIX as a "portfolio" of 30 individual daily implied volatilities, so when a high-volatility day is replaced by a low-volatility one, this term of VIX should drop by $\sigma^2_{d,H} - \sigma^2_{d,L}$.

The 2nd term is more complex, since it depends not only on the day added and the day subtracted, but also on the intervening days. Since the persistence of shocks to expected growth is less than 1, the impact of announcements today on the conditional variance of $\mu_{t+\tau}$ will be smaller than the impact of an announcement later in the next 30 days. In particular, if A_{t+31} is also an announcement day, this 2nd term in VIX could actually increase by a small amount at date $t+1$. However, $A_{t+31} = 1$ and $A_{t+j} = 0$ for $j = 2, \ldots, 30$ maximizes the increase in this 2nd term for any value of ϕ. Furthermore, the 2nd highest value, if A_{t+31} is 0, is negative for any value of ϕ. Thus, provided we assume $A_{t+31} = 0$ for all dates t, the model predicts a drop in VIX from before to after announcements. This assumption, if false, biases against our finding the results we report in the paper.

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