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The Asset-Pricing Implications of Government Economic Policy Uncertainty

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Using the news-based measure of Baker et al. [Baker SR, Bloom N, Davis SJ (2013) Measuring economic policy uncertainty. Working paper, Stanford University, Stanford, CA] to capture economic policy uncertainty (EPU) in the United States, we find that EPU positively forecasts log excess market returns. An increase of one standard deviation in EPU is associated with a 1.5% increase in forecasted three-month abnormal returns (6.1% annualized). Furthermore, innovations in EPU earn a significant negative risk premium in the Fama–French 25 size–momentum portfolios. Among the Fama–French 25 portfolios formed on size and momentum returns, the portfolio with the greatest EPU beta underperforms the portfolio with the lowest EPU beta by 5.53% per annum, controlling for exposure to the Carhart four factors as well as implied and realized volatility. These findings suggest that EPU is an economically important risk factor for equities.

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1. Introduction

Uncertainty about the future has real implications for economic agents' behavior (Bernanke 1983, Bloom 2009, Bloom et al. 2007, Dixit 1989). Government policy makers can contribute to uncertainty regarding fiscal, regulatory, or monetary policy, which we refer to as *economic policy uncertainty* (EPU). Government economic policy is important if for no other reason than a government's own investment and expenditures in the economy; in 2009, for example, federal, state, and local government expenditures in the United States totaled $5.9 trillion, 42.45% of the gross domestic product.[1] Furthermore, the ubiquity of government policy makes it very hard to diversify against. Thus, uncertainty related specifically to the economic policy of governments may impact financial markets.[2]

In this paper we test the impact of economic policy uncertainty on asset prices in the time series and cross section. Economic policy uncertainty helps to forecast log excess returns on the stock market over the two- to three-month horizon beyond other measures of uncertainty and standard controls. We verify that well-known econometric problems associated with persistent forecasters such as EPU do not spuriously drive the forecasting results. Using standard generalized method of moments (GMM) tests, we find that innovations in economic policy uncertainty command a significant negative risk premium in the cross section of stock returns, even when controlling for innovations in other uncertainty measures, in addition to market, size, value, and momentum factors.

We use the Baker et al. (2013) measure of policy uncertainty.[3] The measure is a weighted average of three measures of EPU. The first, with the greatest weight, is the frequency of major news discussing economic policy-related uncertainty. In addition to having the greatest weight, the news-based measure is also the most direct proxy for policy uncertainty in the context of the model of Pástor and Veronesi (2013). In this model the news shocks that drive policy uncertainty are ultimately what affect asset prices above and beyond other economic state variables. The other components of the measure are based on expiring tax provisions and forecaster disagreement about government purchases and inflation. The appeal of the Baker et al. (2013) measure is that it allows for a continuous tracking of policy risk compared with the alternatives.

[1] This figure deducts transfers from the federal to state governments, which would add approximately 10% more to the figure reported (Barnett 2011, U.S. Census Bureau 2012, U.S. Government Printing Office 2014).

[2] Knight (1921) established a distinction between risk and uncertainty. Risk refers to the possibility of a future outcome for which the probabilities of the different possible states of the world are known. Uncertainty refers to a future outcome that has unknown probabilities associated with the different possible states of the world. When referring to EPU we mean uncertainty or risk because we cannot ascertain with any certainty the probabilities of the future direction policy makers will take.

[3] Measures based on news have become a useful way to observe certain phenomenon at a higher frequency than was allowed previously (e.g., Da et al. 2014).

Empiricists frequently measure the impact of policy on asset prices by performing an event study around the policy change. Some studies consider elections as a source of government policy uncertainty (e.g., Belo et al. 2012, Boutchkova et al. 2012, and others). A discrete event has the advantage of being well documented with a timeline of events leading up to the culmination of the event of interest, but it can be artificially precise. For instance, the passing of legislation or an election does not necessarily indicate the complete resolution of uncertainty surrounding the government policy.

The continuous news-based approach we employ has some advantages over an event study. First, news-based EPU measures are available on an ongoing basis. Elections occur infrequently and so only capture short intervals of uncertainty resolution. Second, news-based measures quantify uncertainty resolution rather than assume that a new regime resolves uncertainty.

To guide our investigation of the asset-pricing implications of EPU, we consider the simple framework of the Merton (1973) intertemporal capital asset-pricing model (ICAPM). In the cross section, the ICAPM implies that expected excess returns should vary with conditional exposures to innovations in state variables that forecast investment opportunities, that is, the distribution of future returns on the aggregate wealth portfolio. If EPU adversely affects investment opportunities, we should see a negative relation between the average excess return of an asset and that asset's sensitivity to innovations in EPU. As a first step to determine whether EPU affects investment opportunities, we first investigate whether the EPU measure helps to forecast expected excess stock market returns in the time series. Then, we examine whether exposure to EPU is priced in the cross section of returns.

We find a negative contemporaneous correlation between changes in EPU and market returns. We also find modest evidence of a positive relationship between current levels of EPU and future market excess returns at the two- to three-month horizon, consistent with EPU having real asset-pricing implications. We try to tease out why increases in EPU result in lower prices.

From basic financial theory, the observed decrease in prices with rising policy uncertainty can be due to negative changes in current or expected future cash flows or increases in discount rates. Hence, we investigate whether changes in EPU bear any systematic relationship with current or future dividends. More broadly, theoretical work shows that uncertainty can impact future cash flows including real investment and gross domestic product (GDP) (Aizenman and Marion 1993, Hermes and Lensink 2001, Born and Pfeifer 2014). Empirical work confirms these results (Erb et al. 1996, Hassett and Metcalf 1999, Julio and Yook 2012). We do not find significant evidence of a relationship between changes in EPU and dividend growth for up to two years, with or without controls. The results suggest that the variation in contemporaneous returns related to EPU is driven primarily by variation in expected returns.

We find that EPU commands a risk premium in the cross section of U.S. stock returns. We employ a stochastic discount factor-based GMM technique to estimate the EPU risk premium. We use the Fama–French 25 portfolios formed on size and momentum returns as test assets, although we confirm our results with the 25 size and book-to-market portfolios. We consider eight pricing kernel specifications based on the Carhart (1997) four-factor model plus several subsets of an EPU mimicking portfolio as well as mimicking portfolios for implied volatility and realized volatility. The EPU factor commands a significant negative risk premium with both sets of common factors. Furthermore, the EPU factor commands a significant pricing kernel coefficient in the presence of the other uncertainty measures. Taken together, the evidence suggests that the policy uncertainty factor commands a significant risk premium and helps to price assets above and beyond current models.

This paper is most closely related to the work of Baker et al. (2013) and Pástor and Veronesi (2013), who both associate the Baker, Bloom, and Davis EPU index to various measures of macroeconomic conditions. Using a variety of impulse response functions, Baker et al. (2013) find that positive shocks to their policy uncertainty index are associated with significant decreases in industrial production, employment, GDP, and real investment for at least two to three years. They also associate a number of large swings in the S&P 500 index to policy-related events. Similarly, Pástor and Veronesi (2013) find that the policy uncertainty index negatively correlates with economic conditions, as measured by the Chicago Fed National Activity Index, industrial production growth, the Shiller price-earnings ratio, and the default spread. They also find that stock returns are more volatile and more correlated when policy uncertainty is higher, especially in bad economic times.

Pástor and Veronesi (2013) also relate expected excess returns on the stock market with the Baker, Bloom, and Davis EPU index. Based on the implications of their model, they test whether EPU commands a higher risk premium in bad economic times by relating 3-, 6-, and 12-month future excess market returns to the interactions between the Baker, Bloom, and Davis EPU index and the measures of economic conditions mentioned above. They find modest evidence of such relationships with these specifications.

We incorporate the Pástor and Veronesi (2013) findings into our analysis as well, adding other macro and uncertainty variables as controls. Our time-series analysis differs from that of Pástor and Veronesi (2013) in that we focus on the unconditional relationship between EPU and expected returns.

The rest of this paper is organized as follows. Section 2 reviews the theoretical motivation for the link between EPU and asset prices. Section 3 describes the data and the construction of variables. Section 4 examines the time-series relationship between EPU and stock returns. Section 5 studies the cross-sectional relationship between EPU and stock prices. Section 6 concludes.

2. Theoretical Motivation

In this paper, we test whether EPU increases the equity risk premium and is priced in the cross section of stock returns. We test this hypothesis by using the framework of Merton's (1973) ICAPM under the guidance of the causal mechanisms from the models of Pástor and Veronesi (2012, 2013).

Merton's ICAPM implies the following equilibrium relation between risk and return:

$$\begin{aligned} &E_t(r_{it+1} - r_{ft+1}) \\ &\quad = A \cdot \mathrm{Cov}_t(r_{it+1}, r_{mt+1}) + B \cdot \mathrm{Cov}_t(r_{it+1}, \Delta x_{t+1}), \end{aligned} \tag{1}$$

where r_{ft+1} is the risk-free rate, r_{it+1} is the ex post return on asset i, r_{mt+1} is the ex post market return, and x_t is a vector of state variables that shift the investment opportunity set. The term $\mathrm{Cov}_t(r_{it+1}, r_{mt+1})$ denotes the covariance conditional on information available at time t. The term A is the relative risk aversion of market investors and the term B represents the covariance price of risk for shifts in the state vector that governs the stochastic investment opportunity set x_t. Equation (1) states that in equilibrium, investors are compensated with higher expected returns for bearing systematic market risk and for bearing the risk of unfavorable shifts in the investment opportunity set. That is, investors will have greater demand for assets that hedge against adverse changes in the transition probabilities of future returns on the market portfolio, bidding up their price and driving down their expected returns.

Considerable evidence exists that economic uncertainty is a relevant state variable affecting the investment opportunity set (e.g., Bloom 2009, Bloom et al. 2007, and others). More recently, Pástor and Veronesi (2012, 2013) have presented models suggesting that (economic) policy uncertainty raises the equity risk premium and in particular is also a state variable affecting the investment opportunity set.

A natural question is whether EPU is inherently distinct from general economic uncertainty. That is, does EPU contain any relevant economic information beyond just its contribution to general uncertainty? Pástor and Veronesi (2013) suggest that this is the case. In this model, EPU reflects agents learning about the political costs associated with the implementation of different policies. Agents receive noisy signals (e.g., the news) that change their posterior beliefs about which political forces will get their way, and these signals are driven largely by news shocks that are orthogonal to those driving economic fundamentals. The underlying shocks are ultimately what command the risk premiums, and it is not obvious ex ante that the news shocks driving policy carry the same price of risk as the other shocks driving general economic uncertainty.

To express the point concretely, consider two independent Brownian motion processes, dw_t and dz_t. Suppose that dw_t represents a shock to uncertainty about aggregate real productivity and dz_t represents a pure policy news shock. Suppose furthermore that changes in general economic uncertainty U_t and EPU_t can each be expressed as linear combinations of dw_t and dz_t:

$$\begin{aligned} dU_t &= a \cdot dw_t + b \cdot dz_t, \\ d\,EPU_t &= c \cdot dw_t + d \cdot dz_t. \end{aligned} \tag{2}$$

This suggests that general uncertainty U_t is driven by both sources of uncertainty. Given that economic uncertainty (see Bloom 2009) and policy uncertainty (see §3) both go up during periods of economic weakness, we expect policy uncertainty to be driven by both sources of uncertainty as well. However, we expect that the policy news shock generates a greater portion of the variation in policy uncertainty than in general uncertainty. That is, we assume $d/c > b/a$ so that (a, b) and (c, d) are linearly independent. This allows for the separation of the independent shocks to general uncertainty and EPU given a measure of each. General asset-pricing theory suggests that the processes dw and dz are what actually command risk premiums, and, a priori, we do not know whether each carries the same price of risk. Assuming that the two prices are different, EPU will be an important state variable for asset pricing distinct from general economic uncertainty.

Overall, if the news-based shocks are priced, a good measure of EPU will help to predict expected returns controlling for general uncertainty and economic distress. Hence, we include both such controls in our tests related to the risk premium and EPU. If EPU, like volatility, represents deterioration of the investment opportunity set, then in the language of the ICAPM, EPU should carry a negative price of risk. This would mean that assets with a greater (i.e., less negative) covariance with innovations in EPU should act as a hedge against the deterioration of the investment opportunity set and thus have lower expected returns than assets with a lesser (more negative) covariance with innovations in EPU. Equivalently, a portfolio that is long in assets with the lowest exposure to EPU and short in assets with the greatest exposure to EPU should be compensated with positive expected excess returns.

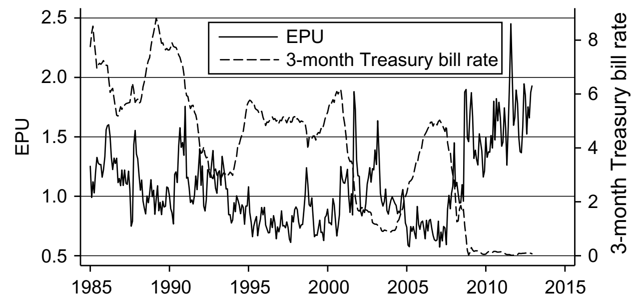
3. Data and Variable Construction

The data in this paper come from a variety of sources. We use the monthly Center for Research in Security Prices (CRSP) value-weighted index as a measure of overall U.S. stock market performance. CRSP generates the index with and without dividends, allowing us to impute a time series of U.S. market-wide dividend growth and dividend–price ratios. Monthly returns, share prices, and numbers of shares outstanding for all common stocks (share code 10 or 11) also come from CRSP. We obtain the Fama–French three factors (*MKT*, *SMB*, and *HML*), the Carhart momentum factor (*UMD*), and the Pástor–Stambaugh liquidity factor (*LIQ*) from Wharton Research Data Services along with the returns on the Fama–French 30 industry portfolios and *VXO*, the Chicago Board Options Exchange monthly index of implied volatility on Standard and Poor’s 100 (S&P 100) index. Bansal and Yaron (2004), Bloom (2009), and Veronesi (1999), among others, link economic uncertainty with stock price volatility. We create a monthly volatility index (*VOL*) by computing the standard deviation of daily stock returns within a month. We create a monthly variance index (*VAR*) as the square of *VOL*. Note that the *VXO* series starts in 1986 but *VAR* and *VOL* go back until 1985.

We also create a variety of business cycle variables from several series available from the Federal Reserve. *BILL* is the yield on the three-month Treasury bill. *RREL* is equal to *BILL* minus the 12-month rolling average of *BILL*. *TERM* is the yield on the 10-year Treasury bond minus the yield on the three-month Treasury bill. *DEFAULT* is the Moody’s BAA corporate bond yield minus the Moody’s AAA corporate bond yield. *CFI* is the Chicago Fed National Activity Index. We also create a smoothed dividend yield D/P given by the 12-month rolling sum of monthly dividends on the CRSP value-weighted index scaled by current Price (see, e.g., Ang and Bekaert 2007).

Our objective is to use a measure that captures the degree of EPU in asset-pricing tests. To the best of our knowledge, only news-based measures such as that of Baker et al. (2013) meet this criterion. Baker et al. (2013) form an index as a weighted average of three distinct components to capture EPU.[4] The first and

Figure 1 **EPU and Treasury Bill Rate**



Note. This figure presents a time-series plot of the Baker, Bloom, and Davis EPU index as well as the yield on the three-month U.S. Treasury bill.

most heavily weighted component measures the percentage of articles in 10 large newspapers relating to economic policy-related uncertainty. The second component is a measure of the magnitude of federal tax code provisions set to expire. The third component measures dispersion of economic forecasts of the consumer price index and purchases of goods and services by federal, state, and local governments. We use the Baker et al. (2013) measure of EPU, denoted *EPU*, for our asset-pricing tests.

The theoretical channel through which EPU affects risk premiums is based on the relative likelihoods that various policies will be adopted. However, as noted by Pástor and Veronesi (2013), the likelihood of any change increases in bad times when economic agents perceive economic policies to be suboptimal. Hence, EPU should be countercyclical. That is, EPU should be high during recessions or other times of economic distress. It should also be higher when general uncertainty is higher as well as when EPU contributes to general uncertainty.

To illustrate the association between EPU and economic distress, Figure 1 presents a plot of the three-month Treasury bill rate and the Baker, Bloom, and Davis EPU index. Times of economic weakness are frequently associated with flights to quality, in which investors flee from riskier assets to Treasury bills, consequently driving down their yields. Consistent with EPU being higher in periods of economic weakness, we see elevated levels of *EPU* coinciding with lower levels in the Treasury bill rate.

More formally, panel A of Table 1 presents the estimates of a standardized linear regression of *EPU* on

[4] We have created our own alternative news-based EPU measure based on the Access World News database and have extended it internationally. We used this alternative measure in earlier versions of the paper. As a result of helpful comments, we now focus just on the U.S. equity market. Consequently, we rely on the Baker, Bloom, and Davis EPU measure because it satisfies all of our needs for a domestic study. We have repeated the tests in this paper with our alternative measure of EPU and have qualitatively the same (in fact, usually stronger) results.

several standard economic state variables (e.g., Santa-Clara and Valkanov 2003, etc.) In a standardized regression the variables are scaled by their standard deviation. The resulting interpretation of the coefficients is that a one-standard-deviation change in variable X results in a beta-standard-deviation change in the dependent variable.

Both measures of uncertainty, the monthly variance of daily returns *VAR*, and the VIX index on the S&P 100 index *VXO* are positively and significantly correlated with *EPU*. The countercyclical term spread *TERM*, smoothed dividend-yield D/P, and default spread *DEFAULT* load significantly and positively on *EPU*. Likewise, the procyclical relative bill rate *RREL* and Chicago Fed National Activity Index *CFI* load significantly and negatively on the index. Combined, these state variables explain approximately 40% of the variation in *EPU* and show that *EPU* is higher in times of economic weakness. However, over half of the variation cannot be explained by these macroeconomic state variables alone.

To investigate the relationship between changes in *EPU* and contemporaneous returns, we estimate the following regression:

$$r_t = \alpha + \beta' \Delta EPU_t + \gamma \Delta X_t + \epsilon_t, \tag{3}$$

where r_t denotes the log excess return of the CRSP value-weighted portfolio in month t, and X_t denotes month t values of the variables in Table 1. The results are reported in panel B of Table 1. Not surprisingly, the first difference of *EPU* significantly and negatively correlates with the excess return on the market. The standard deviation of r_t is $\sigma(r_t) = 4.67\%$. Hence, the

Table 1 Economic Determinants of EPU Index

Panel A: Standardized regressions of EPU on uncertainty and business cycle variables

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VXO	0.41***							0.27***
	(8.08)							(3.75)
VAR		0.34***						−0.01
		(6.59)						(−0.21)
TERM			0.40***					0.27***
			(8.03)					(5.59)
DEFAULT				0.52***				0.20***
				(11.17)				(2.99)
RREL					−0.34***			−0.07
					(−6.62)			(−1.35)
CFI						−0.33***		−0.08
						(−6.51)		(−1.29)
Log(D/P)							0.23***	0.17***
							(4.26)	(3.56)
N	324	336	336	336	336	336	336	324
Adj.-R^2	0.17	0.11	0.16	0.27	0.11	0.11	0.05	0.39

Panel B: Regression of log excess market return on first differences of EPU and controls

	(1)	(2)
Δ *EPU*	−0.10**	−0.28***
	(−2.05)	(−3.63)
Δ *VAR*	0.21**	
	(2.07)	
Δ *VXO*	−0.76***	
	(−13.65)	
Δ *TERM*	−0.10**	
	(−2.39)	
Δ *DEFAULT*	−0.16***	
	(−3.12)	
Δ *RREL*	−0.06	
	(−1.42)	
Δ Log(D/P)	−0.19***	
	(−3.81)	
N	323	335
Adj.-R^2	0.544	0.073

Table 1 (Continued)

Panel C: AR(p) coefficients and DF-GLS statistics

	EPU		*VXO*		*VAR*	
	(1)	(2)	(3)	(4)	(5)	(6)
AR(1)	0.77***	0.71***	0.83***	0.83***	0.57***	0.52***
	(14.11)	(12.74)	(26.36)	(24.19)	(12.52)	(10.52)
AR(2)	−0.02	0.00				
	(−0.24)	(0.04)				
AR(2)	−0.02	0.03				
	(−0.27)	(0.46)				
AR(4)	0.19***	0.18***				
	(3.41)	(3.36)				
MKT_{t-1}		−0.80***		1.14		−5.82**
		(−3.80)		(0.18)		(−2.13)
DF-GLS	−2.29**		−5.29***		−8.11***	
N	332	332	323	323	335	335
Adj.-R^2	0.74	0.75	0.68	0.68	0.32	0.33

Notes. Panel A presents estimates of a standardized linear regression of the Baker, Bloom, and Davis EPU index on a set of economic state variables. *VAR* is the monthly variance of daily returns on the CRSP value-weighted return index. *VXO* is the implied volatility series on the S&P 100 index. *TERM* is the spread between three-month and 10-year U.S. Treasury bond yields. *DEFAULT* is the spread between AAA and BAA bond yields. *RREL* is the yield on the three-month U.S. Treasury bill minus its 12-month rolling average. *CFI* is the Chicago Fed National Activity Index. Log(D/P) is the smoothed log dividend–price ratio on the CRSP value-weighted return series, where D represents the 12-month rolling sum of dividends. Panel C presents AR(p) models of *EPU*, *VXO*, and *VAR* as well as each with lagged market excess returns to isolate innovations for asset-pricing tests. Beneath the estimated coefficients are Dickey–Fuller generalized least squares (DF-GLS) statistics that test the null of a unit root in each time series. Estimations with *VXO* use 324 monthly observations (1986:1–2012:12). Estimations with *VOL* use 336 monthly observations (1985:1–2012:12). For interpreting the effects in panel B, note that $\sigma(r_{t+1}) = 4.67\%$.

** and *** represent significance at the 5% and 1% levels, respectively.

standardized coefficient on ΔEPU of -0.10 (with controls) indicates that a one-standard-deviation value of ΔEPU is associated with a $-0.10 * 4.67\% \approx -47$-basis point log excess return in month t, controlling for changes in the controls. Likewise, without controls, the coefficient of -0.28 indicates that the same one-standard-deviation value of ΔEPU would be associated with a $-0.28 * 4.67\% \approx -1.31\%$ log excess return, not accounting for other variables.

In the next section, we consider *EPU* as a forecaster of stock returns. Given the negative correlation between changes in *EPU* and stock returns, it is necessary to determine how persistent *EPU* is in the time series because the two conditions combined have been shown to bias forecasting tests (Stambaugh 1999, Campbell and Yogo 2006, and others). Furthermore, in §5 we consider *EPU* and the cross section of returns. From the *EPU* time series we must extract innovations, because shocks to risk factors are what command risk premiums. Hence, we consider the time-series properties of *EPU* as well as the other uncertainty measures *VAR* and *VXO* and then isolate innovations in each series. We estimate AR(p) processes for *EPU*, *VXO*, and *VAR*. Following Campbell and Yogo (2006), we use the lag length that minimizes the Bayesian information criterion (BIC; with up to 12 lags). The minimum BIC occurs with one lag for *VAR* and *VXO* and four lags for *EPU*. Hence, we estimate

$$EPU_t = a + b_1 EPU_{t-1} + b_2 EPU_{t-2} + b_3 EPU_{t-3} + b_4 EPU_{t-4} + \epsilon_t^{\mathrm{EPU}}, \tag{4a}$$

$$VXO_t = a + b_1 VXO_{t-1} + \epsilon_t^{\mathrm{VXO}}, \quad \text{and} \tag{4b}$$

$$VAR_t = a + b_1 VAR_{t-1} + \epsilon_t^{\mathrm{VAR}}. \tag{4c}$$

The odd-numbered columns of panel C of Table 1 present these estimates. We also present Dickey–Fuller generalized least squares (DF-GLS) test statistics of the null hypothesis that each series has a unit root. *EPU* is somewhat persistent with the AR(1) coefficient being 0.77. However, we reject at the 5% level the null hypothesis that *EPU* has a unit root.

From the time series we extract the innovations in *EPU* for the asset-pricing tests in §5. In addition to the lags of each variable in Equations (4a)–(4c), we also add one lag of the excess return[5] on the market into the right-hand side of Equations (4a)–(4c). That is, we estimate

$$EPU_t = a + b_1 EPU_{t-1} + b_2 EPU_{t-2} + b_3 EPU_{t-3} + b_4 EPU_{t-4} + c_1 MKT_{t-1} + \epsilon_t^{\mathrm{EPU}}, \tag{5a}$$

[5] We also consider the other state variables from Table 1 but these do not add forecasting power to EPU.

$$VXO_t = a + b_1\, VXO_{t-1} + c_1\, MKT_{t-1} + \epsilon_t^{\mathrm{VXO}}, \tag{5b}$$

$$VAR_t = a + b_1\, VAR_{t-1} + c_1\, MKT_{t-1} + \epsilon_t^{\mathrm{VAR}}. \tag{5c}$$

Then, as innovations, we consider the time series: $\hat{\epsilon}_t^{\mathrm{EPU}}$, and $\hat{\epsilon}_t^{\mathrm{VAR}}$ from Equations (5a) and (5c), respectively, and $\hat{\epsilon}_t^{\mathrm{VXO}}$ from Equation (4b) because the market does not significantly load in Equation (5b). Adding relevant forecasters to Equations (5a)–(5c), such as the lag of the market excess return, helps to isolate the unexpected component of each series.

4. Economic Policy Uncertainty and the Time Series of Stock Returns

In this section we present evidence that EPU forecasts future stock market returns. We also test whether the negative correlation between changes in *EPU* and returns is due to changes in current or expected future cash flows as opposed to discount rates by examining the time-series relationship between dividend growth and *EPU*. We conclude that EPU increases the equity risk premium over time.

4.1. Stock Return Forecasts

To measure the link between expected stock returns and *EPU* we estimate a variety of time-series forecasting regressions of the form

$$r_{t+1,\,t+h} = \alpha + \beta'\, EPU_t + \gamma' X_t + \varepsilon_{t+1}, \tag{6}$$

where r_{t+1} denotes the log excess return on the CRSP value-weighted index during month $t+1$, and $r_{t+1,\,t+h} = r_{t+1} + \cdots + r_{t+h}$ denotes the log excess return on the CRSP value-weighted index during months $t+1$ through $t+h$. That is, $r_{t+1} = \log(1 + R_{m,\,t+1}) - \log(1 + R_{f,\,t+1})$, where $R_{m,\,t+1}$ denotes the holding period return on the CRSP value-weighted index in month $t+1$, and $R_{f,\,t+1}$ denotes the holding period return on the one-month Treasury bill in month $t+1$.

EPU is relatively persistent in that its first autoregressive coefficient is 0.77. Also, Table 1 shows that first differences of *EPU* are negatively correlated with contemporaneous returns. Kendall (1954) and Marriott and Pope (1954) find that high persistence can bias estimates of serial correlations in finite samples. Stambaugh (1999) and Campbell and Yogo (2006), among others, extend the finding to show that these two conditions can also bias predictability regressions in finite samples. The predictability bias is particularly relevant for forecasting with scaled-price variables like dividend yields, whose innovations are highly negatively correlated with returns and have persistence very close to that of a unit root process. The issues associated with persistent variables are less of a concern for *EPU* because the persistence of *EPU* is relatively low compared to scaled-price variables, and *EPU* innovations do not have a large negative correlation with returns (the AR(4) residuals of *EPU* have a correlation of -0.28 with one-month log excess returns).

Nonetheless, we consider the extent to which the forecasting regressions given by Equation (6) suffer from this bias. First, we use Hodrick (1992) standard errors that allow for conditional heteroscedasticity and, more importantly, the error structure induced by the use of overlapping log excess returns as in Equation (6). Hodrick (1992) and Ang and Bekaert (2007) find that, even with persistent regressors such as dividend yields, the standard errors described by Hodrick (1992) maintain good statistical size properties relative to Newey and West (1987) and Hansen and Hodrick (1980) standard errors in that they do not overreject the null of no predictability. In the online appendix (available as supplemental material at http://dx.doi.org/10.1287/mnsc.2014.2044), we consider a series of Monte Carlo experiments and verify that the forecast estimates do not suffer from the substantial size distortions that plague other persistent forecasters in small samples.

Under the null hypothesis that EPU does not affect the equity risk premium, the beta on *EPU* should equal zero in each of the forecasting regressions in Equation (6). Table 2 reports the results of the estimates of Equation (6). The *t*-statistics, based on Hodrick (1992) standard errors, are reported in parentheses below the coefficients.

Panel A of Table 2 presents the estimates of Equation (6) with no controls. The point estimate on *EPU* is positive in all specifications and significant at the 10% level for horizons of two, three, and six months.

In panel B of Table 2, we add the other uncertainty measures, *VAR* and *VXO*, as forecasters. Controlling for these two other uncertainty measures does not significantly change the two-month forecast point estimate significantly. However, the six-month forecast point estimate of *EPU* is no longer significant. The three-month forecasting coefficient on *EPU* just makes the 5% significance cutoff, and the point estimate increases from 4.13 to 4.64. To gauge the economic significance, note that the standard deviation of *EPU* is 0.33. Therefore, a one-standard-deviation increase in *EPU* raises expected excess quarterly returns by $0.33 \cdot 4.64 = 1.53\%$, or 6.12% per annum. As described in §2, *EPU* aims to capture policy news shocks that are correlated with, but distinct from, other uncertainty measures. Therefore, controlling for other measures of uncertainty helps to identify the effect of the policy uncertainty shocks.

Table 2 Forecasting of Log Excess Stock Returns on the Market with Economic Policy Uncertainty

	($h=1$)	($h=2$)	($h=3$)	($h=6$)	($h=12$)
			Panel A: Univariate forecasts		
EPU	0.50	2.82*	4.13*	7.02*	11.11
	(0.53)	(1.69)	(1.85)	(1.82)	(1.64)
N	336	336	336	336	336
Adj.-R^2	−0.00	0.02	0.02	0.04	0.05
			Panel B: EPU forecasts with *VXO* and *VAR*		
EPU	1.02	3.29*	4.64**	6.67	11.34
	(1.13)	(1.93)	(2.00)	(1.60)	(1.46)
VAR	−0.59**	−0.78**	−1.08**	−0.73	−0.10
	(−2.50)	(−2.45)	(−2.41)	(−0.97)	(−0.10)
VXO	0.10*	0.15*	0.22*	0.18	−0.04
	(1.86)	(1.81)	(1.91)	(0.93)	(−0.11)
N	324	324	324	324	324
Adj.-R^2	0.04	0.05	0.06	0.04	0.04
			Panel C: EPU forecasts with controls		
EPU	1.14	4.20*	5.86*	7.68	8.28
	(0.97)	(1.86)	(1.93)	(1.55)	(1.02)
VAR	−0.60**	−0.77**	−1.10**	−0.88	−0.64
	(−2.39)	(−2.37)	(−2.57)	(−1.28)	(−0.71)
VXO	0.14**	0.24***	0.35***	0.43**	0.41
	(2.37)	(2.59)	(2.81)	(2.02)	(1.11)
TERM	−0.19	−0.29	−0.35	−0.00	2.71
	(−0.71)	(−0.55)	(−0.48)	(−0.00)	(1.16)
DEFAULT	−0.23	−1.60	−1.71	−1.30	2.04
	(−0.18)	(−0.69)	(−0.54)	(−0.27)	(0.27)
RREL	0.51	1.29*	2.19**	4.76**	9.14**
	(1.32)	(1.76)	(2.06)	(2.27)	(2.32)
Log(D/P)	1.56*	2.57	3.57	6.39	9.83
	(1.73)	(1.44)	(1.37)	(1.22)	(0.97)
N	324	324	324	324	324
Adj.-R^2	0.05	0.08	0.12	0.14	0.23

Notes. Each column of panels A, B, and C represents the estimated coefficients from a regression of the form

$$r_{t+1,t+h} = \alpha + \beta' EPU_t + \gamma' X_t + \epsilon_{t+1},$$

where $r_{t+1,t+h}$ denotes the excess log return on the market during month $t+1$ through $t+h$; $h = 1, 2, 3, 6$, or 12; *EPU* denotes the Baker, Bloom, and Davis EPU index; and X denotes a set of controls that includes *VAR*, the variance of daily returns on the CRSP value-weighted return during month t; *VXO*, the implied volatility series on the S&P 100 index; *TERM*, the difference between the yields on the 10-year Treasury bond and the three-month Treasury bill; *DEFAULT*, the difference between the yields on BAA and AAA corporate bonds; *RREL*, the yield on the three-month Treasury bill minus its 12-month rolling average; and Log(D/P), the smoothed dividend–price ratio on the CRSP value-weighted index. t-Statistics based on Hodrick (1992) heteroscedasticity and serial correlation-robust standard errors are given in parentheses in panels A, B, and C.

*, **, and *** represent significance at the 10%, 5%, and 1% levels, respectively.

In panel C of Table 2, we add four standard control variables to the analysis (see e.g., Cooper and Priestly 2009, Lettau and Ludvigson 2001, and others). The additional control variables help to ensure that *EPU* is directly impacting returns and that it is not acting as an instrument for other forecasters such as dividend yields or interest rates. The two- and three-month forecasting slopes of *EPU* are still positive and significant at the 10% level. Overall, the predictive results from Table 2 indicate a forecasting role for EPU at the two- to three-month horizon.

4.2. Dividend Growth

Panel B of Table 1 shows a negative correlation between current market returns and changes in *EPU*. Such a relationship suggests that changes in *EPU* are negatively correlated with current or expected future dividend growth on the market index, or positively

Table 3 Economic Policy Uncertainty and Log Dividend Growth

	1 month	3 months	6 months	12 months	24 months
	Panel A: Log dividend growth on EPU (no controls)				
Δ *EPU*	−0.18	−0.08	−0.03	0.00	−0.02
	(−1.34)	(−0.40)	(−0.15)	(0.01)	(−0.08)
N	335	335	335	335	324
Adj.-R^2	0.00	0.00	−0.00	−0.00	−0.00
	Panel B: Log dividend growth on EPU (with controls)				
Δ *EPU*	−0.11	−0.13	−0.05	0.03	−0.00
	(−0.57)	(−0.62)	(−0.21)	(0.11)	(−0.02)
Δ *VAR*	−0.01	0.01	0.01	−0.00	0.00
	(−0.65)	(0.65)	(0.53)	(−0.15)	(0.17)
Δ *VXO*	0.00	−0.00	−0.00	−0.00	−0.00
	(0.17)	(−0.02)	(−0.19)	(−0.24)	(−0.12)
Δ *TERM*	−0.14	−0.05	−0.09	−0.10	−0.11
	(−0.83)	(−0.27)	(−0.29)	(−0.21)	(−0.16)
Δ *DEFAULT*	−0.12	−0.16	−0.04	−0.17	−0.14
	(−0.36)	(−0.38)	(−0.05)	(−0.16)	(−0.13)
Δ *RREL*	0.19	−0.01	0.05	−0.04	0.11
	(0.79)	(−0.02)	(0.11)	(−0.06)	(0.14)
N	323	323	323	323	312
Adj.-R^2	0.01	0.02	0.02	0.04	0.03

Notes. Panel A presents estimates of the regression of the log-dividend growth of the CRSP value-weighted index over 1–24 months on the first difference of the Baker, Bloom, and Davis EPU index. Panel B presents estimates of similar regressions but with the addition of the business-cycle control variables. The control variables include *VAR*, the variance of daily returns on the CRSP value-weighted return during month t; *VXO*, the implied volatility series on the S&P 100 index; *TERM*, the difference between the yields on the 10-year Treasury bond and the three-month Treasury bill; *DEFAULT*, the difference between the yields on BAA and AAA corporate bonds; *RREL*, the yield on the three-month Treasury bill minus its 12-month rolling average; and Log(D/P), the smoothed dividend–price ratio on the CRSP value-weighted index. t-Statistics based on Hodrick (1992) standard errors are given in parentheses.

correlated with changes in discount rates. In this section, we test the null hypothesis that changes in *EPU* do not affect current or future dividend growth. To obtain log dividend growth $\Delta d_{t,\,t+m}$ over months $t, \ldots, m-1$, we add the one-month log dividend growth ($\Delta d_t = \log(D_t) - \log(D_{t-1})$):

$$\Delta d_{t,\,t+m} = \sum_{i=1}^{m} \Delta d_{t+i-1}. \tag{7}$$

To test the null hypothesis that changes in *EPU* do not affect dividend growth, we estimate the following regression:

$$\Delta d_{t,\,t+m} = \alpha + \beta' \Delta EPU_t + \gamma' X_t + \varepsilon_{t+1}. \tag{8}$$

The results are in Table 3. Panel A of Table 3 presents estimates of the linear regression in Equation (8) with no control variables. Column 1 shows the 1-month results; column 2 shows the 3-month results; and columns 3, 4, and 5 show the 6-, 12-, and 24-month dividend growth results, respectively. For all columns, changes in *EPU* are statistically insignificant. Panel B of Table 3 repeats the analysis but includes the full set of control variables used previously. Regardless of the time horizon, the *EPU* coefficient is statistically insignificant. We fail to find evidence that *EPU* affects dividend growth. Hence, it is more likely that price drops associated with increases in *EPU* are due to increases in expected returns, resulting in the same expected dividends discounted at a higher rate.

5. Economic Policy Uncertainty and the Cross Section of Returns

In this section, we investigate the cross-sectional implications of EPU being a state variable that affects the investment opportunity set. That is, we test whether exposure to *EPU* is priced in the cross section of stock returns.

As discussed in §2, theoretical work suggests that EPU may command a risk premium observable in the cross section of stock returns. Pástor and Veronesi (2012) model firms with differing exposure to policy uncertainty. They posit that firms with higher exposure to policy uncertainty typically have higher expected returns, although the phenomenon is state dependent and can potentially have the opposite effect. Likewise, innovations in policy uncertainty adversely affect investment opportunities by increasing uncertainty.

5.1. Factor-Mimicking Portfolios

Linear factor models such as the ICAPM (1) imply linear pricing kernel models of the form:

$$E[m_t R_t^e] = 0, \tag{9}$$

$$m_t = a + b'(f_t - E(f)), \tag{10}$$

where R_t^e denotes a vector of excess returns, f_t is a vector of factors, a is a constant and b is a vector of coefficients (see, e.g., Cochrane 1996). Consider the projection $b'R_t^e$ of the pricing kernel m_t onto the space of excess returns, R_t^e. That is,

$$m_t = b'R_t^e + \epsilon_t, \tag{11}$$

$$E(\epsilon_t R_t^e) = 0. \tag{12}$$

From Equations (9)–(12), it follows that

$$0 = E(m_t R_t^e) = E((b'R_t^e + \epsilon_t)R_t^e) = E((b'R_t^e)R_t^e). \tag{13}$$

The projection, called the *factor-mimicking portfolio* of m, is just a regression of the discount factor on the returns R_t^e. From Equation (13), we see that the mimicking portfolio is also a discount factor that prices assets in R_t^e. In particular, $b'R_t^e$ contains all the relevant asset-pricing information that m_t does but does not have the portion ϵ_t, which is uninformative for pricing R_t^e; ϵ_t may, in principle, also include measurement error that is orthogonal to the asset returns. Hence, it is convenient in empirical work to form mimicking portfolios for a proposed discount factor because they can reduce measurement error and filter out the information that is irrelevant to the prices of the test assets.

By the linearity of the projection operator and of the discount factor (Equation (10)), the factor-mimicking portfolio of the discount factor can be obtained from a linear combination of the mimicking portfolios for each individual factor. Following Breeden et al. (1989), Ang et al. (2006), and others, we create factor-mimicking portfolios F_X for $X = EPU, VXO$, or VAR, by estimating the following regression over the entire sample period:

$$\hat{\epsilon}_t^X = a_X + b_X' R_t^e + \eta_t^X, \quad X = EPU, VXO, \text{ or } VAR, \tag{14}$$

where $\hat{\epsilon}_t^X$ denotes the innovations for variable $X = EPU$, VXO, or VAR, estimated in (5a), (4b), and (5c), respectively, and R_t^e denotes the excess returns on a set of basis assets. Specifically, the mimicking portfolios are given by

$$F_X = \hat{b}_X' R_t^e, \quad X = EPU, VXO, \text{ or } VAR, \tag{15}$$

where $\hat{b}_X$ is the estimate from Equation (14). We choose the excess returns on the Fama–French 25 size and momentum portfolios. These portfolios produce a large spread in average monthly returns over our sample period (from −6 to 134 basis points per month), and these returns are not explained by only a few factors like the average returns on the size and book-to-market portfolios (see, e.g., Lewellen et al. 2010).

Table 4 presents estimates of regressions of the factor-mimicking portfolios on the innovations $\hat{\epsilon}_t^X$, $X = EPU$, VAR, or VXO as well as the differences

Table 4 Properties of Factor-Mimicking Portfolios

	F_{EPU}	F_{EPU}	F_{EPU}	F_{VAR}	F_{VAR}	F_{VAR}	F_{VXO}	F_{VXO}	F_{VXO}
$\hat{\epsilon}^{EPU}$	0.14***	0.07***	0.07***		0.54*	0.48		0.87	1.03
	(7.44)	(4.04)	(3.83)		(1.70)	(1.49)		(1.00)	(1.17)
$\hat{\epsilon}^{VAR}$		0.00*	0.00	0.31***	0.12***	0.09***		0.01	−0.08
		(1.92)	(1.00)	(12.13)	(4.09)	(2.97)		(0.15)	(−0.98)
$\hat{\epsilon}^{VXO}$		0.01***	0.01***		0.12***	0.12***	0.54***	0.53***	0.54***
		(7.33)	(7.56)		(8.51)	(8.81)	(19.60)	(14.14)	(14.59)
Δ *TERM*			0.01			0.23			0.57
			(1.09)			(1.13)			(1.03)
Δ *DEFAULT*			0.05*			0.91*			3.64***
			(1.84)			(1.96)			(2.89)
Δ *RREL*			−0.01			−0.24			0.48
			(−0.69)			(−0.91)			(0.67)
Δ Log(D/P)			0.18			5.00**			19.21***
			(1.23)			(2.00)			(2.82)
N	332	323	323	335	323	323	323	323	323
Adj.-R^2	0.14	0.37	0.38	0.30	0.45	0.46	0.54	0.54	0.56

Notes. This table presents estimates from several ordinary least squares regressions of the factor-mimicking portfolios on the innovations in *EPU*, the uncertainty measures, and the control variables. The control variables include *TERM*, the difference between the yields on the 10-year Treasury bond and the three-month Treasury bill; *DEFAULT*, the difference between the yields on BAA and AAA corporate bonds; *RREL*, the yield on the three-month Treasury bill minus its 12-month rolling average; and Log(D/P), the smoothed dividend–price ratio on the CRSP value-weighted index. t-Statistics are given in parentheses.
*, **, and *** represent significance at the 10%, 5%, and 1% levels, respectively.

of the other control variables from Table 1, that is, regressions of the form

$$F_{Xt} = a + b_1\hat{\epsilon}_t^{\text{EPU}} + b_2\hat{\epsilon}_t^{\text{VAR}} + b_3\hat{\epsilon}_t^{\text{VXO}} + c'Z_t + \eta_t,$$
$$X = EPU,\ VXO,\ \text{or}\ VAR. \tag{16}$$

Each mimicking portfolio correlates significantly with its corresponding innovation. Furthermore, the innovations in *VXO* correlate significantly with both the *VAR* and *EPU* factor-mimicking portfolios. This is not surprising given the correlation between *EPU* and uncertainty and implied and realized volatility. The results indicate that the tests including both F_{EPU} and F_{VXO} will be the most informative about isolating the asset-pricing implications of EPU.

5.2. GMM Estimation of Factor Risk Premiums

We use standard GMM tests of the linear factor model in Equations (9) and (10). We estimate the risk premiums associated with *EPU* and determine whether the *EPU* factor-mimicking portfolio F_{EPU} is a factor that helps to price assets. As test assets we use the Fama–French 25 size and momentum portfolios. This set of test assets lacks the strong three-factor structure of the Fama–French 25 portfolios formed on size and book-to-market ratio identified by Lewellen et al. (2010). We follow the two-step GMM method given by Cochrane (2005) and used by Brennan et al. (2004), among others. We test the asset-pricing equation

$$E[m_t R_{pt}^e] = 0, \tag{17}$$

with the linear pricing kernel

$$m_t = 1 + b'(f_t - \bar{f}), \tag{18}$$

where b is a vector of constants and $f_t - \bar{f}$ is a demeaned set of factors from the Fama–French three-factor (FF3F) model, the Carhart momentum factor, the Pástor–Stambaugh liquidity factor, the *EPU* mimicking factor F_{EPU}, the *VXO* mimicking factor F_{VXO}, and the *VAR* mimicking factor F_{VAR}. We choose the common normalization $E(m_t) = 1$ because the mean of the discount factor is not identified when exclusively using excess returns as test assets.

Given a matrix W, a GMM estimate, $\hat{b}(W)$, of b minimizes the quadratic form $g_T(b)'Wg_T(b)$ where $g_T(b) = T^{-1}\sum_{t=1}^{T} m_t(b)R_t^e$ and $m_t(b)$ is given by Equation (18). The one-step GMM estimator is simply $\hat{b}(I)$, where I denotes the identity matrix. The one-step estimator effectively treats the pricing of all assets in R_t^e as equally important. The two-step GMM estimator is given by $\hat{b}(\hat{S}^{-1}(\hat{b}(I)))$, where S is the estimated covariance matrix of the pricing errors from the one-step estimation. That is, $\hat{S}$ is an estimate of

$$S = \sum_{j=-\infty}^{\infty} E(u_t u_{t-j}'), \tag{19}$$

where $u_t = m_t R_t^e$. Following Vassalou (2003), Cochrane (1996), and others, we use the Newey–West (1987) estimator of S:

$$\hat{S} = T^{-1}\sum_{t=1}^{T}\sum_{j=-k}^{k}\left(\frac{k-|j|}{k}\right)u_t u_{t-j}'. \tag{20}$$

The Newey–West estimator of S is typically used because it accounts for possible serial correlation and is positive definite in every sample. We report estimates with $k = 4$, but they are virtually identical to those estimated without allowing for serial correlation ($k = 0$).

Panel A of Table 5 presents the estimates of the b values and panel B presents the estimated risk premiums and Hansen–Jagannathan distances (Hansen and Jagannathan 1997). Column (1) of panel A demonstrates that the excess return on the market is a significant factor to help to price assets. The negative coefficients on *SMB* and *HML* suggest that they covary negatively with marginal utility over this period, which results in their carrying a positive risk premium (see panel B). However, the pricing kernel coefficient on *SMB* has no significance unless F_{VAR} is included. The pricing kernel coefficient on *HML* is significant at the 1% level in all specifications. Momentum also loads significantly and negatively in the pricing kernel in every specification. Column (2) adds F_{EPU}. The F_{EPU} coefficient is 1.37 and is statistically insignificant taken by itself; however, it carries a significantly negative risk premium (panel B), suggesting that investors will demand lower returns to hold assets that hedge against increases in *EPU*, that is, assets with positive *EPU* betas. Note that an increase in EPU represents a deterioration of the investment opportunity set and thus should correlate positively with marginal utility (hence with m_t), resulting in a negative risk premium.

Column (3) of Table 5 uses F_{VXO} instead of F_{EPU} and shows it also has a statistically insignificant discount factor coefficient of -0.09 taken alone but a significantly negative risk premium (panel B). This is consistent with the evidence from Ang et al. (2006) that *VXO* has a negative price of risk. Finally, taking F_{EPU} and F_{VXO} together in column (4) reveals a significant pricing kernel coefficient for each. Taken together with the evidence from Table 4, and the theory described in §2, it is likely that the presence of F_{VXO} helps to separate the policy uncertainty news shocks in F_{EPU} from the other uncertainty shocks captured by F_{VXO}. Column (5) uses the other uncertainty factor F_{VAR} and column (6) adds F_{EPU}. Like F_{VXO}, F_{VAR} does not have a significant discount factor coefficient unless it is considered with F_{EPU}. However, once considered with *EPU*, the discount factor coefficient is negative and significant, whereas the coefficient on F_{EPU} is positive and significant; F_{VAR}, however, does not have a significant risk premium when considered with F_{EPU}.

Table 5 **GMM Estimations of Linear Asset-Pricing Models with Economic Policy Uncertainty**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				Panel A: Pricing kernel coefficients				
MKT	−7.03***	−5.77***	−13.21***	−18.41***	−10.36***	−12.65***	−17.72***	−6.41***
	(−4.34)	(−2.95)	(−3.16)	(−4.08)	(−4.11)	(−4.63)	(−3.31)	(−2.79)
SMB	−1.36	−0.73	−2.85	−2.86	−3.59	−7.99***	−7.58***	0.91
	(−0.71)	(−0.36)	(−1.36)	(−1.33)	(−1.52)	(−3.05)	(−2.91)	(0.42)
HML	−10.77***	−9.85***	−13.21***	−14.12***	−13.94***	−20.26***	−19.70***	−10.85***
	(−3.31)	(−2.97)	(−3.8)	(−4.2)	(−3.59)	(−4.77)	(−4.77)	(−3.08)
UMD	−3.57***	−3.54**	−4.18***	−5.07***	−5.25***	−9.35***	−8.25***	−5.15***
	(−2.59)	(−2.53)	(−2.93)	(−3.38)	(−3.21)	(−4.65)	(−3.96)	(−3.15)
LIQ								12.76**
								(2.21)
F_{EPU}		1.37		5.01***		8.79***	8.06***	2.51*
		(1.02)		(3.05)		(4.57)	(4.20)	(1.80)
F_{VXO}			−0.09	−0.22***			−0.12	
			(−1.58)	(−3.31)			(−1.18)	
F_{VAR}					−0.19	−0.78***	−0.57***	
					(−1.64)	(−4.71)	(−2.59)	
				Panel B: Estimated risk premiums (in % per month)				
MKT	0.96***	0.97***	0.90***	0.90***	0.85***	0.52*	0.56*	1.17***
	(3.39)	(3.39)	(3.18)	(3.05)	(2.98)	(1.87)	(1.87)	(3.44)
SMB	0.05	0.07	0.02	0.03	0.01	0.07	0.05	0.03
	(0.27)	(0.37)	(0.1)	(0.16)	(0.07)	(0.41)	(0.29)	(0.14)
HML	0.60**	0.55**	0.64***	0.58***	0.65***	0.74***	0.70***	0.69***
	(2.51)	(2.35)	(2.80)	(2.75)	(2.72)	(3.25)	(3.10)	(2.86)
UMD	0.33	0.35	0.39	0.55*	0.44	0.64**	0.60*	0.57*
	(1.15)	(1.19)	(1.33)	(1.80)	(1.50)	(2.02)	(1.88)	(1.83)
LIQ								−1.87**
								(−2.24)
F_{EPU}		−1.35***		−1.40***		−1.03**	−0.96**	−1.52***
		(−3.22)		(−3.23)		(−2.48)	(−2.26)	(−3.40)
F_{VXO}			−70.50***	−66.07***			−34.56	
			(−2.91)	(−2.73)			(−1.51)	
F_{VAR}					−21.40**	−11.35	−11.22	
					(−2.46)	(−1.47)	(−1.40)	
HJ distance	0.41	0.40	0.40	0.38	0.39	0.40	0.40	0.41
Std. err.	0.05	0.05	0.05	0.05	0.05	0.06	0.06	0.06

Notes. This table presents estimates from a two-step GMM estimation of several linear asset-pricing models. The test assets are the excess returns on the Fama–French 25 portfolios formed on size and momentum return (from month $t-12$ through month $t-2$). Each column represents a pricing kernel that is a different linear combination of the possible factors. The list of factors consists of the Fama–French three factors, the Carhart momentum factor, the Pástor–Stambaugh traded liquidity factor, the *EPU* factor-mimicking portfolio F_{EPU}, the *VXO* factor-mimicking portfolio F_{VXO}, and the *VAR* factor-mimicking portfolio F_{VAR}. Panel A presents estimates of the coefficients on the pricing kernel, and panel B presents estimated risk premiums, in percent per month, associated with each factor, along with the Hansen–Jagannathan (HJ) distance for each model. Results are based on 332 time-series observations of each variable (1985:5–2012:12). t-Statistics are in parentheses and based on Newey–West standard errors with four lags.

*, **, and *** represent significance at the 10%, 5%, and 1% levels, respectively.

Column (7) of Table 5 considers F_{EPU}, F_{VXO}, and F_{VAR} along with the Carhart four-factor model; F_{EPU} and F_{VAR} have a significant discount factor coefficient, but F_{VXO} does not, indicating that *VXO* may not add any marginal asset-pricing information after *VAR*, *EPU*, and the Carhart four factors are considered. Furthermore, only F_{EPU} has a significant risk premium after controlling for F_{VAR} and F_{VXO}. Finally, column (8) of Table 5 adds the liquidity factor of Pástor and Stambaugh (2003). This loads significantly but does not appear to explain the relationship between F_{EPU} and returns.

Panel B of Table 5 contains the estimated risk premiums for each factor. The market has an estimated risk premium ranging from 0.52% to 1.17% per month. *SMB* and *HML* have positive risk premiums, but only *HML* is statistically significant in the presence of any other factor. Momentum's risk premium is only significant in the presence of an uncertainty factor. The liquidity factor carries a statistically significant risk premium of −1.87%. The *EPU* factor-mimicking portfolio F_{EPU} carries a statistically significant risk premium ranging from −0.96% to −1.52% in each specification.

We report the Hansen–Jagannathan distances for each column in panel B of Table 5. Typically, asset-pricing models are compared by observing the Hansen–Jagannathan distance, and the lowest value is deemed the best performer. The Hansen–Jagannathan distance does not vary significantly across specifications. Although we cannot meaningfully identify a best-performing model, the results consistently show that *EPU* is a priced risk factor.

To gauge the economic significance of the estimates in Table 5, we determine the expected returns in each of the 25 size- and momentum-sorted portfolios that can be attributed to F_{EPU} exposure. In particular, for each of the 25 portfolios, we estimate the portfolio's exposure to the factors by estimating the regression

$$R^e_{it} = a + \beta_{i,\text{MKT}} MKT_t + \beta_{i,\text{SMB}} SMB_t + \beta_{i,\text{HML}} HML_t + \beta_{i,\text{UMD}} UMD_t + \beta_{i,F_{\text{EPU}}} F_{\text{EPU},t} + \beta_{i,F_{\text{VXO}}} F_{\text{VXO},t} + \beta_{i,F_{\text{VAR}}} F_{\text{VAR},t} + \eta_t, \quad (21)$$

where $i = Small$, 2, 3, 4, Big and $j = Low$, 2, 3, 4, $High$. Table 6 reports the estimates of $\beta_{i,F_{\text{EPU}}}$, the estimated exposure to F_{EPU} controlling for all the other factors.

The return R_{12} has the highest $\beta_{i,F_{\text{EPU}}}$, 0.18, among the 25 size- and momentum-sorted portfolios, and R_{51} has the lowest $\beta_{i,F_{\text{EPU}}}$ of -0.30. The spread between these two F_{EPU} betas is 0.48. From column (7) of Table 5, the F_{EPU} risk premium is -0.01 when controlling for the other six factors in Equation (21), so the lowest *EPU*-beta portfolio has a return that is $-0.48 * (-0.96\%) * 12 = 5.53\%$ lower per annum than that of the highest *EPU*-beta portfolio. This represents a plausible but nontrivial premium to hedge against exposure to EPU.

To economize on space, we do not show all 25 of the analogous F_{VXO} and F_{VAR} betas. The spread between the highest and lowest $\beta_{i,F_{\text{VXO}}}$ across the 25 portfolios is 0.02. From column (7) of Table 5, the risk premium on F_{VXO} is -34.56% per month so the lowest F_{VXO}-beta portfolio has a return that is $-0.02 * (-34.56\%) * 12 = 8.29\%$ lower per annum than that of the highest F_{VXO}-beta portfolio. Similarly, the spread between the highest and lowest $\beta_{i,F_{\text{VAR}}}$ across the 25 portfolios is 0.05. From column (7) of Table 5, the risk premium on F_{VXO} is -11.22% per month so the lowest F_{VAR}-beta portfolio has a return that is $-0.05 * (-11.22\%) * 12 = 6.73\%$ lower per annum than that of the highest F_{VXO}-beta portfolio. Note, however, that the risk premium on F_{VAR} is not statistically significant, so the 6.73% risk contribution may be imprecise. Overall, F_{EPU} contributes the least to risk among the 25 size- and momentum-sorted portfolios, but the contribution is considerable and on the same order of magnitude as that of F_{VXO} and F_{VAR}.

5.3. Robustness Checks

Table 7 presents several alternative GMM specifications to check the robustness of the results in the previous section. The mimicking portfolios for *EPU*, *VXO*, and *VAR* are formed on the sample period 1985:5–2012:12. Column (1) of Table 7 presents estimates of the GMM model with all three uncertainty measures but using one-step GMM. One-step GMM is a useful robustness check for GMM estimations (see, e.g., Cochrane 1996). The one-step procedure equally weights pricing errors, whereas two-step GMM gives more weight to the most statistically informative (that is, the most precisely estimated) sample moments, in estimating model parameters. This in turn prevents the estimation from only pricing combinations of portfolios that are meaningful only in a statistical sense. The one-step estimator will therefore have less power and higher standard errors. Nonetheless, the one-step estimation yields a discount factor coefficient on F_{EPU} that exhibits a negative risk premium that is statistically significant at the 1% level.

Table 6 F_{EPU} **Beta Estimates for 25 Size- and Momentum-Sorted Portfolios: From April 1985 to December 2012**

	$\beta_{i,F_{\text{EPU}}}$					$t(\beta_{i,F_{\text{EPU}}})$				
	Momentum return					Momentum return				
Size	Low	2	3	4	High	Low	2	3	4	High
Small	0.14	0.18	0.15	0.09	−0.10	3.11	7.67	6.09	3.25	−2.78
2	−0.03	−0.04	0.04	0.03	0.02	−1.07	−1.48	1.41	1.15	0.71
3	−0.07	−0.10	0.12	0.01	−0.06	−1.86	−3.57	5.15	0.39	−2.54
4	−0.19	−0.08	0.03	−0.08	−0.12	−5.01	−3.00	1.18	−3.17	−4.61
Big	−0.30	0.10	0.17	0.09	−0.08	−7.24	3.51	8.04	4.39	−3.18

Notes. This table reports estimates of $\beta_{i,F_{\text{EPU}}}$ from the specification in column (7) of Table 5 for the 25 size- and momentum-sorted portfolios. That is,

$$R^e_{it} = a + \beta_{i,\text{MKT}} MKT_t + \beta_{i,\text{SMB}} SMB_t + \beta_{i,\text{HML}} HML_t + \beta_{i,\text{UMD}} UMD_t + \beta_{i,F_{\text{EPU}}} F_{\text{EPU},t} + \beta_{i,F_{\text{VXO}}} F_{\text{VXO},t} + \beta_{i,F_{\text{VAR}}} F_{\text{VAR},t} + \eta_t,$$

where $i = Small$, 2, 3, 4, Big and $j = Low$, 2, 3, 4, $High$. The sample period is from May 1985 to December 2012. Corresponding t statistics are in the panel to the right.

Table 7 Robustness Checks of GMM Estimations of Linear Asset-Pricing Models with Economic Policy Uncertainty

	(1)	(2)	(3)	(4)		(5)
			Panel A: Pricing kernel coefficients			
MKT	−11.49	−11.55***	−6.32*	−18.81	*MKT*	−8.23
	(−1.63)	(−4.85)	(−1.72)	(−1.5)		(−1.62)
SMB	−10.16***	−1.37	1.93	10.65**	*SMB*	0.62
	(−3.18)	(−0.57)	(0.67)	(2.39)		(0.3)
HML	−27.04***	−17.44***	−14.32***	−3.08	*HML*	−6.21***
	(−4.85)	(−4.68)	(−3.6)	(−0.61)		(−2.63)
UMD	−12.08***	−19.49***	−18.48***	−7.03	*UMD*	−2.61*
	(−4.33)	(−5.94)	(−5.54)	(−1.35)		(−1.72)
F_{EPU}	9.51***		5.29*	0.64	$\hat{\epsilon}^{\mathrm{EPU}}$	4.56**
	(3.00)		(1.69)	(0.14)		(2.48)
F_{VXO}	0.106			−0.51*	$\hat{\epsilon}^{\mathrm{VXO}}$	−0.12
	(0.74)			(−1.82)		(−1.25)
F_{VAR}	−1.07***			1.47**	$\hat{\epsilon}^{\mathrm{VAR}}$	−0.04
	(−3.49)			(2.46)		(−0.29)
			Panel B: Estimated risk premiums (% per month)			
MKT	0.59	1.03**	0.98**	1.29***	*MKT*	0.36
	(1.64)	(2.21)	(2.10)	(2.66)		(1.09)
SMB	0.04	0.09	0.04	−0.08	*SMB*	−0.14
	(0.19)	(0.36)	(0.15)	(−0.29)		(−0.70)
HML	1.12***	0.69**	0.62**	0.40	*HML*	0.33*
	(3.61)	(2.44)	(2.26)	(1.55)		(1.79)
UMD	0.76*	3.68***	3.54***	2.69***	*UMD*	0.28
	(1.88)	(5.15)	(4.95)	(3.46)		(0.92)
F_{EPU}	−1.10**		−2.27***	−2.82***	$\hat{\epsilon}^{\mathrm{EPU}}$	−9.87**
	(−2.06)		(−2.96)	(−3.32)		(−2.24)
F_{VXO}	−51.72*			−129.16***	$\hat{\epsilon}^{\mathrm{VXO}}$	52.79
	(−1.67)			(−3.06)		(0.72)
F_{VAR}	−14.55			−77.21***	$\hat{\epsilon}^{\mathrm{VAR}}$	12.83
	(−1.42)			(−4.63)		(0.41)
HJ distance		0.46	0.45	0.45		0.41
Std. err.		0.06	0.06	0.07		0.06

Notes. This table presents GMM estimates of several linear asset-pricing models. Each column represents a pricing kernel that is a different linear combination of the possible factors. Unless otherwise stated the estimates come from the two-step efficient GMM procedure and the sample is 1985:5–2012:12. In columns (1) and (5) the test assets are the excess returns on the Fama–French 25 portfolios formed on size and momentum. In column (1) the estimation is done via one-step GMM. In columns (2), (3), and (4) the test assets are the union of the Fama–French 25 size and book-to-market portfolios and five industry portfolios. In column (5) the sample is 1986:2–2012:12 and the test assets are the union of the Fama–French 25 size- and momentum-sorted portfolios and the four tradable factors. The list of factors consists of the Fama–French three factors, the Carhart momentum factor, the *EPU* factor-mimicking portfolio F_{EPU}, the *VXO* factor-mimicking portfolio F_{VXO}, and the *VAR* factor-mimicking portfolio F_{VAR}. Column (5) uses the innovations $\hat{\epsilon}^{\mathrm{EPU}}$, $\hat{\epsilon}^{\mathrm{VXO}}$, and $\hat{\epsilon}^{\mathrm{VAR}}$ as opposed to their factor-mimicking portfolios. Panel A presents estimates of the coefficients on the pricing kernel and panel B presents estimated risk premiums, in percent per month, associated with each factor along with the Hansen–Jagannathan (HJ) distance for each model. t-Statistics are in parentheses and based on Newey–West standard errors with four lags.

*, **, and *** represent significance at the 10%, 5%, and 1% levels, respectively.

In columns (2), (3), and (4) of Table 7 we keep the factors F_{EPU}, F_{VXO}, and F_{VAR} unchanged but consider another set of test assets, the union of the Fama–French 25 size and book-to-market portfolios with the Fama–French five industry portfolios. The discount factor coefficient on F_{EPU} is insignificant for this set of test assets, but the risk premium is statistically significant at the 1% level. The results are effectively unchanged by removing the five industry portfolios but we include them to obviate the econometric concerns expressed by Lewellen et al. (2010).

Finally, column (5) of Table 7 presents the estimates from the main test but using the innovations in the factors $\hat{\epsilon}^X$, $X = EPU, VAR, VXO$ as opposed to the factor-mimicking portfolios. *EPU* loads significantly in the discount factor and commands a significantly negative risk premium, although, as expected, the factor-mimicking portfolio results in statistically more precise estimates.

6. Conclusion

Government economic policy—including that of taxation, expenditure, monetary, and regulatory—has large, market-wide economic effects that are largely nondiversifiable. Economic agents make real economic decisions based on expectations about the future economic policy environment. Thus, even market-benevolent policy makers can increase risk by generating an environment of uncertainty about their future economic policy decisions.

This paper uses the Baker et al. (2013) measure to examine the role that economic policy uncertainty (EPU) plays in asset prices. EPU is positively correlated with (but distinct from) general economic uncertainty as proxied by the volatility of market returns. An increase of one standard deviation in EPU is associated with a contemporaneous 1.31% decrease in market returns and a 1.53% increase in future three-month log excess returns (6.12% annualized). Changes in EPU do not appear to affect cash flows. That is, changes in EPU have no statistically discernible effect on dividend growth over the 1- to 24-month horizons.

Through a variety of tests we find that EPU matters for the discount rate. Although Pástor and Veronesi (2012, 2013) emphasize the role of EPU in different economic states, we show that the effect exists in general. The cross section and the time series provide evidence that suggests that EPU is an economically important risk factor.

Supplemental Material

Supplemental material to this paper is available at http://dx.doi.org/10.1287/mnsc.2014.2044.

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