Does Variance Risk Have Two Prices? Evidence from the Equity and Option Markets*

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March 02, 2015

Abstract

We formally compare the conditional Variance Risk Premia (VRPs) in the equity and option markets. Both VRPs follow common patterns and respond similarly to changes in volatility and economic conditions. However, we reject the null hypothesis that they are identical and find that their difference is strongly related to measures of the financial standing of intermediaries. These results shed new light on the information content of the option VRP, suggest the presence of market frictions between the two markets, and are consistent with the key role played by intermediaries in setting option prices.

Keywords : Variance Risk Premium, Options, Stocks, Financial Intermediaries

JEL Classification : G12, G13, C58

*We would like to thank Torben Andersen, Turan Bali, David Bates, Geert Bekaert, Sebastien Betermier, Peter Christoffersen, Stefano Giglio, Ruslan Goyenko, Nicolas Mougeot, Elisa Ossola, Lasse Pedersen, Christopher Polk, Sergei Sarkissian, Olivier Scaillet, Viktor Todorov, Fabio Trojani, Andrea Vedolin, Hao Zhou and seminar participants at the Copenhagen Business School, the Federal Reserve Board, the LSE Systemic Risk Centre, the University of Sherbrooke, the 2014 CIRPÉE Applied Financial Time Series Conference, the 2014 Dauphine Amundi Chair in Asset Management Workshop, the 2014 IFSID Conference on Derivatives, the 2014 meeting of the Institute for Mathematical Finance, the 2014 NBER Summer Institute, and the 2014 Paris Finance Conference. The authors gratefully acknowledge financial support from the Montreal Institute of Structured Finance and Derivatives (IFSID). The views expressed in this paper are those of the authors and do not necessarily reflect those of the Bank for International Settlements.

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1 Introduction

The Variance Risk Premium (VRP) is the compensation investors are willing to pay for assets that perform well when stock market volatility is high. Because a version of the conditional VRP can be easily computed from the prices of index options (the option VRP), it is often viewed by academics and policymakers alike as the most readily available proxy for fluctuations in investors' risk aversion and aggregate discount rates.[1] The widespread use of the option VRP implicitly relies on the assumption that risk is priced consistently across markets. However, previous studies provide evidence of potential mispricing between equity and option markets and stress the key role played by financial intermediaries (broker-dealers) in determining option prices.[2] If option prices are driven by local demand and supply forces, the option VRP may behave quite differently from the premium paid by equity investors for stocks that hedge market volatility shocks (the equity VRP).

This paper addresses this issue by formally testing whether the VRPs formed in the equity and option markets are identical. By construction, if the two VRPs are equal, so are their linear projections on any set of predictive variables. Building on this insight, we compare the two markets based on the projections of their VRPs on a rich set of predictors that capture volatility and economic conditions, as well as the financial standing of broker-dealers. The linear framework developed here has several advantages. First, it yields simple expressions for both projections that can be estimated using standard regression techniques. Second, it guarantees that the two premia are fully comparable because they are conditioned on the same set of predictors. Third, it allows us to measure the role played by each predictor in explaining the two VRPs and their potential difference.

[1] See Bali and Zhou (2014), Bekaert and Hoerova (2014), Bollerslev, Gibson, and Zhou (2011), and Drechsler and Yaron (2011), among others.

[2] The relative mispricing of SP500 index options is documented by Constantinides, Czerwonko, Jackwerth, and Perrakis (2011) and Constantinides, Jackwerth, and Perrakis (2009), whereas the role of intermediaries in setting option prices is discussed by Adrian and Shin (2010), Bates (2003, 2008), Chen, Joslin, and Ni (2013), Garleanu, Pedersen, and Poteshman (2009), and Fournier (2014).

Finally, it is consistent with the extensive literature that use linear regressions to measure the conditional market risk premium and forecast realized variance.[3]

The procedure for estimating the VRP projections requires as only inputs a set of variance risk-sensitive portfolios in each market. In the option market, Britten-Jones and Neuberger (2000) show how to form a portfolio that replicates the variance payoff. The price of this portfolio, measured by the squared VIX index, is sufficient to estimate the option-based projection. In the equity market, we follow Ang, Hodrick, Xing, and Zhang (2006) and construct 25 portfolios sorted based on variance and market betas. Then, we specify a two-factor representation of these portfolio returns that includes the variance and market factors, and compute the equity-based projection using an extension of the two-pass regression developed by Gagliardini, Ossola, and Scaillet (2014).

Our empirical results shed new light on the commonalities and differences between the two VRP projections measured at the quarterly frequency. We find that their average levels are both approximately equal to -2.00% per year, consistent with the notion that investors are willing to pay a premium to hedge against volatility shocks. In addition, the magnitude of both premia tends to increase after volatility shocks and during recession periods. However, the empirical evidence formally rejects the null hypothesis that the premia are identical. The difference between the two VRP projections exhibits several key features. First, it changes signs as the option VRP can be either below or above its equity counterpart. Second, it is economically large—in 9 quarters out of 84, its magnitude is above 3.00% per year, which is 1.5 times the average premium itself. Third, it is not exclusively associated with crisis episodes such as the great recession in 2007-08. Finally, its variations are driven by two measures of the financial standing of intermediaries commonly used in the literature, namely the leverage ratio of broker-dealers and the quarterly return of the Prime Broker Index (PBI).[4] We observe that when these

[3] See Fama and French (1989), Keim and Stambaugh (1986), Ferson and Harvey (1991), as well as Campbell, Giglio, Polk, and Turley (2013) and Paye (2012).

[4] See, for instance, Adrian and Shin (2010, 2013) who demonstrate empirically that the leverage ratio drops when intermediaries hit their risk contraints, and Boyson, Stahel, and Stulz (2010) who use the PBI return in the context of hedge fund contagion.

intermediaries deleverage or suffer short-term losses, the magnitude of the option VRP rises significantly, whereas the equity VRP remains unchanged.

Before examining the implications of these results, it is important to verify that they are not simply caused by misspecification—if the two-factor model does not fully capture the returns of the 25 portfolios, the equity-based projection could be biased and lead us to the wrong conclusion that the two markets differ. To address this issue, we first perform formal specification tests of the two-factor model. In particular, we assess the magnitude of the pricing errors and determine whether a linear combination of the equity portfolios can replicate the variance payoff. Second, we study the potential impact of distinguishing between different components of variance risk (e.g., high- versus low-frequency variations). Finally, we explicitly model the time-variation in portfolio betas and include several additional risk factors. Based on this extensive analysis, we are confident that misspecification is not the driving force behind the difference between the two markets.

The rejection of the null hypothesis that the two conditional VRPs are equal yields two main insights. First, it leads to a more nuanced view of the information content of the readily available option VRP. Despite the commonalities between the two markets, caution should be exercised when the option VRP replaces the equity VRP as a measure of the risk aversion of equity investors. For instance, periods in which financial intermediaries deleverage or suffer losses are associated with a significant increase in the magnitude of the option VRP relative to its equity counterpart. Such phenomena are observed in 1998 or in 2008 as a consequence of the collapse of the Long Term Capital Management (LTCM) fund and the recent financial crisis. The opposite situation is also observed during the monetary easing period in early 2000s when the option VRP largely underestimates the premium paid by equity investors to hedge variance risk. Second, the VRP difference indicates the presence of market frictions that prevent the law of one price to apply.

One friction-based explanation borrows from the international finance literature which commonly attributes mispricing across countries to market segmentation (e.g. Bekaert, Harvey, Lundblad, and Siegel (2011)). Theoretically, Basak and Croitoru (2000) demonstrate that pricing discrepancies can exist in equilibrium if investors face portfolio constraints that limit risk sharing. In practice, such constraints may arise because equity investors face information costs or regulatory constraints that limit their positions in the option market or because broker-dealers do not have the mandate to trade in stocks exposed to variance risk. An alternative explanation proposed by Garleanu and Pedersen (2011) is that the price of identical assets can diverge in equilibrium if they are traded in markets with different margin requirements, which is the case for the equity and option markets. While the marginal contribution of each theory cannot be determined without knowing all the constraints faced by investors, our empirical evidence suggests that the margin-based explanation, if used alone, cannot fully account for the path followed by the VRP difference. First, it predicts that the VRP difference should decline (rise) when investors' funding liquidity is high (low). Our results reveal that direct measures of funding liquidity such as the default and TED spreads are weakly related to the VRP difference. Second, this theory cannot easily explain that the VRP difference takes both positive and negative values because the spread in margins is unlikely to change signs.

Another insight from our empirical analysis is the strong relationship between the leverage ratio of broker-dealers and the option VRP, which resonates with the key role played by these institutions in the option market. As shown empirically by Garleanu, Pedersen, and Poteshman (2009) and Chen, Joslin, and Ni (2013), broker-dealers supply options to public investors in exchange for a premium for holding residual risk. Therefore, the extent to which they are able to perform this task should depend on their ability to bear risk and take on leverage—if the latter declines (increases), the option supply should drop (rise) and leads to a higher (lower) price of variance risk in the option market. To test this hypothesis, we examine whether an increase in leverage results in lower option prices. Consistent with a supply-based mechanism, we document a strong and negative relationship between leverage and the VIX index. This evidence is not overturned when we treating leverage as endogenous and controlling for the potential impact of additional predictors.

In addition to the main results presented above, we conduct a detailed analysis of the predictors that track volatility and economic conditions, namely the lagged realized variance, the Price/Earnings (PE) ratio, the default spread, and the quarterly employment and inflation rates. In both equity and option markets, the price of variance risk increases dramatically after volatility shocks. During these volatile periods, investors revise their expectations of future variance upward ("physical expectation" effect), but are also willing to pay a higher price for assets that provide insurance against future volatility shocks ("risk-neutral expectation" effect). We find that the second effect dominates the first and explains why the magnitude of the equity VRP increases. Second, these two effects offset one another for both the default spread and the PE ratio. Therefore, these variables do not affect the premium variation despite they are strong predictors of the future realized variance (e.g., Campbell, Giglio, Polk, and Turley (2013)). Third, lower employment and inflation rates increase the magnitude of the VRP, which helps to explain why the price of variance risk tends to rise during NBER recessions.

Our work is related to several strands of the literature. First, there is an extensive literature on the role played by variance risk in the equity market. Ang, Hodrick, Xing, and Zhang (2006) infer the unconditional VRP from the returns of portfolios exposed to volatility shocks, while Bansal, Kiku, Shaliastovich, and Yaron (2013) and Campbell, Giglio, Polk, and Turley (2013) derive an intertemporal CAPM with stochastic volatility to explain the cross-section of average stock returns. Relative to these papers, we perform a dynamic analysis that allows us to estimate the entire path of the equity VRP and determine the drivers of its time-variation. Second, several studies examine the evolution of the VRP using option prices (e.g., Bollerslev, Gibson, and Zhou (2011), Todorov (2010)). Our dynamic comparison with the equity market sheds new light on the informational content of the option VRP. Third, Constantinides, Jackwerth, and Perrakis (2009) and Constantinides, Czerwonko, Jackwerth, and Perrakis (2011) document violations of stochastic dominance bounds by call and put options written on the SP500 index. We provide a possible explanation for this mispricing, namely the difference in the pricing of variance risk. Finally, Adrian and Shin (2010) and Chen, Joslin, and Ni (2013) show empirically that the behavior of financial intermediaries is an important driver of option prices. Relative to these papers, we find that these intermediaries affect the price of variance risk very differently in the equity and option markets.

The remainder of the paper is organized as follows. Section 2 presents the methodology used to formally comparing the conditional VRPs in the equity and option markets. Section 3 describes the data. Section 4 contains the main empirical findings. Section 5 provides several interpretations for our main findings. Section 6 summarizes the results of the sensitivity analysis, and Section 7 concludes. The appendix provides a detailed description of the estimation procedure and reports additional results.

2 The Empirical Framework

2.1 Comparing the Equity and Option Markets

To begin the presentation of our empirical framework, we define the conditional Variance Risk Premium (VRP) as

$$\lambda_{v,t} = E\left(rv_{t+1}|\, I_t\right) - E^Q\left(rv_{t+1}|\, I_t\right), \tag{1}$$

where rv_{t+1} is the realized variance of market returns between time t and $t+1$, and $E\left(rv_{t+1}|\, I_t\right)$, $E^Q\left(rv_{t+1}|\, I_t\right)$ denote the physical and risk-neutral expectations of rv_{t+1} conditioned on all available information at time t.[5] In frictionless markets, the variance risk has to be priced consistently across the equity and option markets. To test this null hypothesis, we develop a simple comparison approach that is not based on the VRP

[5] This formulation is commonly used in the option literature (e.g., Bollerslev, Tauchen, and Zhou (2009)) and implies a negative value for the VRP if investors wish to hedge against increases in aggregate volatility.

itself, but on its linear projection on the space spanned by predictive variables that track the evolution of volatility and economic conditions, as well as the financial standing of intermediaries:

$$\lambda_{v,t}(z) = proj(\lambda_{v,t}|\, z_t) = F_v' z_t - V_v' z_t, \tag{2}$$

where the J-vector z_t includes a constant and $J-1$ centered predictors, and $F_v' z_t$, $V_v' z_t$ denote the linear projections of the physical and risk-neutral expectations of rv_{t+1}, respectively. By construction, if the VRPs are the same in both markets, so are their linear projections—therefore, any differences between projections signal periods when the price of variance risk differs across markets.[6]

Building on this insight, we compute the two equity- and option-based estimates of $\lambda_{v,t}(z)$ as

$$\begin{aligned} \hat{\lambda}^e_{v,t}(z) &= (\hat{F}_v - \hat{V}^e_v)' z_t, \\ \hat{\lambda}^o_{v,t}(z) &= (\hat{F}_v - \hat{V}^o_v)' z_t, \end{aligned} \tag{3}$$

where the vector $\hat{F}_v$ is obtained from a time-series regression of rv_{t+1} on z_t, and the risk-neutral vectors $\hat{V}^e_v$ and $\hat{V}^o_v$ are extracted from equity and options prices using the estimation procedure described below. To compare both markets, we simply take the difference between the two estimated projections:

$$\hat{D}_t(z) = \hat{\lambda}^e_{v,t}(z) - \hat{\lambda}^o_{v,t}(z) = (\hat{V}^o_v - \hat{V}^e_v)' z_t. \tag{4}$$

The linear framework used here has several notable features. First, it yields simple expressions for the VRP projections and their difference—in particular, $\hat{D}_t(z)$ only depends on $\hat{V}^e_v$ and $\hat{V}^o_v$ as the physical expectation term $\hat{F}_v z_t$ cancels out. Second, it guarantees that the two markets are fully comparable because both VRP projections are conditioned

[6]Note that the opposite does not hold, i.e., the projections can be equal even if the VRPs differ. This situation occurs when the difference between the two VRPs is orthogonal to the predictors.

on the same information set. Third, it allows us to measure the impact of the predictive variables on the price of variance risk in each market.

2.2 Estimation Procedure

The estimation procedure uses as inputs the returns of a set of equity and option portfolios. These portfolios are exposed to variance risk and allow us to make inferences about the risk neutral vectors V_v^e and V_v^o. For sake of brevity, we describe the main steps of the procedure below and relegate in the appendix additional details on the properties of the different estimators, which are all consistent and asymptotically normally distributed.

2.2.1 The Option-Based Vector V_v^o

We begin with the option market and consider a portfolio of options whose excess return r_{t+1}^o replicates the payoff of a variance swap, i.e.,

$$r_{t+1}^o = rv_{t+1} - p_{rv,t}, \tag{5}$$

where $p_{rv,t}$ is the forward price of the realized variance, also called the implied variance iv_t. Britten-Jones and Neuberger (2000) show how to construct such a portfolio, while Carr and Wu (2009) demonstrate that the squared VIX index computed from option prices can be used as a model-free measure of iv_t.[7] Denoting by Q_o the risk-neutral measure formed in the option market, we can write the forward price $p_{rv,t}$ as $E^{Q_o}\left(rv_{t+1} | I_t \right)$, which implies that

$$proj(p_{rv,t}| z_t) = proj(iv_t| z_t) = V_v^{o\prime} z_t. \tag{6}$$

The estimation procedure is straightforward because iv_t is observable at each time t using the squared VIX index.[8] Based on equation (6), we can therefore estimate V_v^o from a

[7] Specifically, a variance swap contract can be replicated with a static portfolio of options that ensures a constant dollar gamma (unit beta to the variance factor) and with positions in market futures to maintain delta-neutrality (zero beta to the market factor).

[8] As shown by Carr and Wu (2009) and Jiang and Tian (2005), the equality between iv_t and the squared VIX only holds approximately in case of large market movements. In the sensitivity analysis,

time-series regression of iv_t on z_t. The only challenge stems from data limitations: whereas rv_{t+1} and z_t are observed over a long period beginning in 1970 (the long sample), iv_t is only available in the early 1990's (the short sample). Therefore, we use the Generalized Method of Moments (GMM) for samples of unequal lengths developed by Lynch and Wachter (2013) to improve the precision of the estimated vector $\hat{V}_v^o$. The basic idea is to adjust the initial estimate of V_v using information about rv_{t+1} and z_t over the long sample. The intuition behind this adjustment can be easily illustrated with the following example. Suppose that we wish to estimate the averages of the realized and implied variances, denoted by rv and iv (i.e., z_t equals 1). Now suppose that the estimated mean of rv_{t+1} over the short sample, denoted $\widehat{rv}_S$, is above the more precise estimate computed over the long sample. Because rv_{t+1} and iv_t are positively correlated, $\widehat{iv}_S$ is also likely to be above average. Therefore, $\widehat{iv}_S$ is adjusted downward to produce the final estimate $\widehat{iv}$.

2.2.2 The Equity-Based Vector V_v^e

For the equity market, we construct a set of 25 portfolios by sorting stocks into quintiles based on their betas on the variance and market factors. These value-weighted portfolios are rebalanced each month to maintain stable exposures to both risk factors. Our portfolio formation essentially mimics the one used by Ang, Hodrick, Xing, and Zhang (2006) to infer the average level of the VRP in the equity market and is described in the appendix. We specify a two-factor model (variance and market) to capture the return of each variance portfolio $r_{p,t+1}^e$ $(p = 1, ..., 25)$. Projecting $r_{p,t+1}^e$ on the space spanned by z_t and the factors yields the following expression:

$$r_{p,t+1}^e = -proj(p_{p,t}|\, z_t) + b_{pv} \cdot rv_{t+1} + b_{pm} \cdot f_{m,t+1} + e_{p,t+1}, \tag{7}$$

where $f_{m,t+1}$ is the market excess return, b_{pv}, b_{pm} denote the constant conditional betas, $e_{p,t+1}$ is the idiosyncratic term, and $p_{p,t}$ is a weighted sum of the forward prices of the two

we re-estimate the vector V_v^o using the SVIX index that is robust to jumps (see Martin (2013)) and document similar results.

risk factors.[9] In equilibrium, $p_{p,t}$ is equal to $b_{pv}E^{Q_e}\left(rv_{t+1}|\,I_t\right)+b_{pm}\cdot E^{Q_e}\left(f_{m,t+1}|\,I_t\right)$, where Q_e is the risk-neutral measure formed in the equity market.[10] Therefore, the projected price $proj(p_{p,t}|\,z_t)$ can be written as

$$proj(p_{p,t}|\,z_t) = c_p' z_t = (b_{pv}\cdot V_v^{e\prime} + b_{pm}\cdot V_m^{e\prime})z_t, \tag{8}$$

where $V_m^{e\prime}z_t$ denotes the projection of the risk-neutral expectation of the market factor. The procedure for estimating V_v^e builds on recent work by Gagliardini, Ossola, and Scaillet (2014) and is a simple extension of the traditional two-pass regression to the conditional setting examined here.[11] In the first step, we run a time-series regression of $r_{p,t+1}^e$ on z_t, rv_{t+1}, and $f_{m,t+1}$ to estimate c_p, b_{pv}, and b_{pm} for each variance portfolio (equation (7)). In the second step, we exploit the restriction that the vector c_p is equal to a linear combination of the two vectors V_m^e and V_v^e (equation (8)) — by running a cross-sectional regression of each element of the estimated vector $\hat{c}_p$ on the estimated betas $\hat{b}_{pm}$ and $\hat{b}_{pv}$, we can therefore estimate each element of V_v^e.

The two-factor representation with constant betas in equation (7) is motivated by the fact that the equity portfolios (i) are sorted along the market and variance dimensions and (ii) are frequently rebalanced to maintain stable exposures to both factors. However, it raises the concern that the estimation procedure imposes more structure on the equity market than on the option market—if V_v^o is equal to V_v^e but the two-factor model is misspecified, $\hat{V}_v^e$ could be biased and lead us to the wrong conclusion that the two VRPs differ. To address this issue, we perform an extensive specification analysis of the equity

[9] Equation (7) can be derived from the excess return definition, $r_{p,t+1}^e = \left(b_p' f_{t+1} + \epsilon_{j,t+1}\right) - p_{p,t}$, where $b_p = [b_{pv}, b_{pm}]'$, $f_{t+1} = [rv_{t+1}, f_{m,t+1}]'$. The last term $p_{p,t}$ equals $R_{f,t}P_{p,t}$, where $R_{f,t}$ is the gross risk-free rate and $P_{p,t}$ is the time-t price of $b_p' f_{t+1}$ conditioned on all available information I_t. Alternatively, $p_{p,t}$ can be written as $b_p' p_{f,t}$, where $p_{f,t} = [R_{f,t}P_{rv,t}, R_{f,t}P_{f_m,t}]'$ is the vector of forward prices of the risk factors.

[10] This equality is equivalent to the one imposed on conditional expected returns, i.e., $E(r_{p,t+1}^e|\,I_t) = b_{pv}\lambda_{v,t}^e + b_{pm}\lambda_{m,t}^e$, where $\lambda_{v,t}^e$, $\lambda_{m,t}^e$ denote the conditional variance and market risk premia.

[11] Equations (7) and (8) are simply the conditional counterparts of the traditional two-pass regression used in an unconditional setting: (i) the time-series regression becomes $r_{p,t+1}^e = -c_p + b_{pv}\cdot rv_{t+1} + b_{pm}\cdot f_{m,t+1} + e_{p,t+1}$, where c_p is a scalar; (ii) the cross-sectional regression becomes $c_p = b_{pv}\cdot V_v^e + b_{pm}\cdot V_m^e$, where V_v^e and V_m^e are the unconditional risk-neutral expectations (i.e., $E^{Q_e}(rv_{t+1}) = V_v^e$, $E^{Q_e}(f_{m,t+1}) = V_m^e$).

VRP. First, we conduct several formal specification tests of the two-factor model and measure the ability of the equity portfolios to replicate the variance payoff. Second, we study whether grouping the different variance components (e.g., low- and high-frequency) into a single factor can potentially bias the empirical results. Third, we explicitly allow for time-variation in portfolio betas. Fourth, we include additional sources of risk such as the size, book-to-market, and momentum factors. This analysis presented in the empirical section allows us to rule out equity-based misspecification as the driving force behind the empirical results.

3 Data Description

3.1 Predictive Variables

We conduct our empirical analysis using quarterly data between April 1970 and December 2012. We employ a set of five macro-finance predictors to capture volatility and economic conditions: the lagged realized variance, the Price/Earnings (PE) ratio, the quarterly inflation rate, the quarterly growth in aggregate employment, and the default spread (all of which are expressed in log form). The theoretical motivation for using these variables as well as their ability to predict realized variance are discussed in the recent studies of Bollerslev, Gibson, and Zhou (2011), Campbell, Giglio, Polk, and Turley (2013), and Paye (2012).[12] The PE ratio is downloaded from Robert Shiller's webpage and is defined as the price of the SP500 divided by the 10-year trailing moving average of aggregate earnings. Inflation data are computed from the Producer Price Index (PPI), aggregate employment is measured by the total number of employees in the nonfarm sector (seasonally-adjusted), and the default spread is defined as the yield differential between Moody's BAA- and AAA-rated bonds. These three series are downloaded from the Federal Reserve Bank of St. Louis.

[12] In the appendix, we also examine the role played by additional macro-finance variables such as the dividend yield, the growth rate of industrial production, and the term spread.

In addition to the macro-finance variables mentioned above, we consider two measures of the financial standing of broker-dealers (both expressed in log form). The first measures the leverage of broker-dealers using data from the Federal Reserve Flow of Funds Accounts (Table L 128).[13] While we use the leverage ratio defined as the asset to equity value as our main measure, we also perform the estimation using the annual change in leverage. Adrian and Shin (2010, 2013) provide supporting evidence that broker-dealers actively manage their leverage levels based on their risk-bearing capacity—in good times, they increase their leverage and expand their asset base, whereas they deleverage in bad times, possibly because of tighter Value-at-Risk constraints or higher risk aversion levels. Second, we borrow from Boyson, Stahel, and Stulz (2010) and compute the value-weighted index of publicly-traded prime broker firms, including Goldman Sachs, Morgan Stanley, Bear Stearns, UBS, and Citigroup. The quarterly return of this index allows us to capture changes in the financial strength of the major players in the brokerage sector.

Table 1 provides summary statistics for the predictors. To facilitate comparisons across the estimated coefficients presented in the empirical section, all predictors are standardized.[14] The comparison of the persistence levels for the two broker-dealer variables reveals that they contain information at different frequencies. The leverage ratio is a slow-moving predictor that proxies for long-term changes in the risk-bearing capacity of financial intermediaries, whereas the PBI return captures the short-term reaction of these intermediaries to aggregate losses. Perhaps unsurprisingly, the two broker-dealer variables also capture some business cycle fluctuations — for instance, the correlation between the leverage and PE ratios equals 0.33. To explicitly distinguish between the two sets of predictors, we therefore regress the leverage ratio and the PBI return on the macro-finance variables and take the residual components from these regressions.

[TABLE 1 HERE]

[13] The Federal Reserve defines broker-dealers as financial institutions that buy and sell securities for a fee, hold an inventory of securities for resale, or both.

[14] Lettau and VanNieuwerburgh (2008), among others, provide empirical evidence that the mean of financial ratios exhibit substantial structural shifts after 1991. Therefore, we follow their recommendation and allow for the possibility that predictors have different means before and after 1991.

3.2 Equity Portfolios

As discussed in the previous section, we construct a set of 25 portfolios to extract information about the VRP in the equity market. To summarize the properties of these portfolios, we take an equally-weighted average of all portfolios in the same variance beta quintile (Low, 2, 3, 4, High). Panel A of Table 2 shows the relationship between average returns and post-formation variance betas over the period 1970-2012. For each portfolio, the variance beta is obtained from the two-factor model in equation (7), where the market and variance factors are proxied by the quarterly excess return of the CRSP index and the quarterly sum of the daily squared SP500 returns, respectively. In a multi-period setting, risk-averse investors wish to hedge against increases in aggregate volatility because such changes represent a deterioration in investment opportunities. Therefore, stocks that perform poorly in periods of high volatility should command higher expected returns (e.g., Campbell, Giglio, Polk, and Turley (2013)). Consistent with this view, we find that the low variance portfolio loads negatively on the variance factor (with a post-ranking beta of -0.61) and yields an average return of 7.42% per year. As we move toward the high variance portfolios, the post-ranking beta increases by 0.76 and the average return drop by 2.99% per year.

Next, Panel B examines whether commonly-used asset pricing models explain the average return difference across portfolios. Whereas high volatility shocks are associated with stock market declines (the correlation between factor innovations equals -0.49), the two factors capture different dimensions of risk because the CAPM alphas exhibit the same pattern as the average portfolio returns. For the Fama-French model, the alphas remain different from zero, which may not be surprising given that the portfolios have similar size and book-to-market levels (see Panel A). When we include traded momentum and Pastor-Stambaugh liquidity factors, the models still fail to capture the cross-section of average returns.[15]

[15] As shown in the appendix, the short sample (1992-2012) unveils similar patterns for average returns, post-formation betas, and alphas.

Finally, and as a prelude to our formal specification tests, we measure the response of the variance portfolio returns to extreme variance shocks. We find that during the three largest volatility shocks (Oct. 2008, Oct. 1987, July 2011), the market-hedged return of the high minus low variance portfolios is always positive and takes an average value as large as 8.66% per quarter. For the quarters with the lowest variance shocks (Jan. 2012, July 2000, Jan. 1975), we observe the exact opposite pattern with a quarterly average return of -3.00%. All of these results provide supportive evidence that the returns of the equity portfolios are exposed to variance risk and can be used to extract information regarding its premium.

[TABLE 2 HERE]

4 Main Empirical Results

The empirical section of the paper contains three parts. First, we compute the linear projection of the VRP in each market (equity, option) and determine how it relates to changes in the macro-finance and broker-dealer variables. Second, we measure the difference between the VRP projections to identify when the two markets diverge. Third, we examine whether such divergences are potentially explained by the misspecification of the two-factor representation of equity portfolio returns.

4.1 Equity and Option Variance Risk Premia

4.1.1 Realized Variance Predictability

To begin our empirical analysis, we report in Table 3 the vector $\hat{F}_v$ obtained from the predictive regression of the realized variance on the predictors—as shown in equation (3), this vector is a required input for measuring the equity and option VRP projections. Panel A contains the estimated coefficients associated with the (standardized) macro-finance variables. In the first row, the lagged realized variance is used as the sole variable in the predictive regression and produces a strongly positive coefficient that captures the persistent component of the variance process. In the second row, we condition on all macro-finance variables simultaneously. There is a positive and statistically significant relationship between the default spread and the future realized variance. A natural explanation for this result is that risky bonds are short the option to default. When the expected future variance is above average, investors bid down the price of risky bonds, which in turn increases the default spread. Conditional on the other predictors, a high PE ratio also signals above-average future variance and helps to capture episodes during which both stock prices and volatility are high. All of these results are in line with the evidence documented by Campbell, Giglio, Polk, and Turley (2013) and Paye (2012) over the same quarterly frequency.

Building on previous work by Brunnermeier and Pedersen (2009), Paye (2012) suggests that financial intermediation could amplify shocks to asset markets in periods when financial intermediaries experience deleveraging spirals. Contrary to this view, Panel B reveals that the incremental power of the broker-dealer variables is weak because none of the t-statistics is significantly different from zero.

[TABLE 3 HERE]

4.1.2 Equity Market

Next, we compute the estimated vector $\hat{F}_v - \hat{V}_v^e$ that drive the time-variation of the equity VRP projection. While $\hat{F}_v$ is taken from Table 3, the risk-neutral vector V_v^e is estimated using the conditional two-pass regression described in Section 2. Panel A of Table 4 presents the coefficients associated with the macro-finance variables and reveals that the average equity VRP equals -2.24% per year $(-0.58 \cdot 4)$. The lagged realized variance has a significant impact on the VRP, both statistically and economically, i.e., a one-standard deviation increase in realized variance increases the magnitude of the VRP projection by 1.36% per year (-0.34 $\cdot$ 4). The intuition for this result is simple: in volatile periods, assets that pay off when future volatility increases further becomes extremely valuable and this effect dominates the increase in expected future variance documented in Table 3 (i.e., $\hat{V}_v^{e\prime} z_t > \hat{F}_v^{\prime} z_t$). Next, we observe that the physical and risk-neutral expectation effects offset one another for both the PE ratio and the default spread because the estimated coefficients are not statistically significant. Therefore, these variables have little impact on the equity VRP despite the fact that they strongly forecast realized variance (as shown in Table 3 and in previous studies). Finally, the coefficients associated with the inflation and employment rates are both positive. As both predictors tend to be high during expansions, they help capture the countercyclical component of the equity VRP. However, only past inflation exhibits a statistically significant coefficient.

The resulting VRP projection, computed as $(\hat{F}_v - \hat{V}_v^e)' z_t$, is plotted in Figure 1 over the period 1970-2012. Its value is negative for most quarters, consistent with the notion that investors are willing to pay a premium for stocks that perform well when volatility increases. With an autocorrelation coefficient of 0.44, it also inherits some of the persistence exhibited by the predictors. The premium is characterized by transitory spikes following large volatility shocks such as the 1987 crash, the burst of the dotcom bubble, or the 2008 crisis. Finally, it is generally countercyclical, as illustrated by the 1973-74 and 2008-09 recessions.

Turning to the analysis of the broker-dealer variables, we find in Panel B that their relationships with the equity VRP are weak. The coefficients associated with the leverage ratio, the change in leverage, or the PBI return are all close to zero and their t-statistics far below the conventional significance thresholds. Therefore, measures of the financial standing of financial intermediaries have little influence on the pricing of variance risk in the equity market.[16]

[TABLE 4 HERE]

[FIGURE 1 HERE]

[16] The appendix also reveals that the paths of the VRP projections computed with and without the broker-dealer variables are nearly indistinguishable.

4.1.3 Option Market

Repeating the analysis for the option market, we compute the vector $\hat{F}_v - \hat{V}_v^o$, where the risk-neutral vector $\hat{V}_v^o$ is obtained by regressing the implied variance on the predictors using the GMM procedure described in Section 2. The implied variance is measured as the squared VIX index computed from three-month SP500 option prices and available over the short sample (1992-2012).[17] For the macro-finance variables, the coefficients reported in Panel A of Table 5 are comparable with those estimated in the equity market, except for the PE ratio which exhibits a positive and significant coefficient. The overall evidence therefore suggests that the equity and option VRPs respond similarly to volatility and business cycle conditions.

A more striking result documented in Panel B is the strong and positive relationships between the two broker-dealer variables and the option VRP projection. Periods when broker-dealers deleverage or suffer short-term losses are associated with a higher magnitude for the option VRP, whereas the opposite holds when their leverage or stock returns are above average. The estimated coefficient for the leverage ratio is not only highly significant, it is also economically large, i.e., a one-standard deviation decrease in leverage increases the magnitude of the premium by 1.40% per year (-0.35·4). Because the two orthogonalized broker-dealer variables are negatively correlated (-0.26), the predictive information contained in the PBI return is obscured when used alone in the regression. Adding the leverage ratio clarifies the relationship between the PBI return and the option VRP and produces a positive and statistically significant coefficient (0.18). Finally, the rightmost columns of Panel B confirm that all of these results remain unchanged when the leverage ratio is replaced with the annual change in leverage.

To visualize these findings, we plot in Figure 2 the time-variation of the option VRP projection measured as $(\hat{F}_v - \hat{V}_v^o)' z_t$, where z_t includes all predictors.[18] For comparison

[17]The quarterly VIX index is also referred to as the VXV index and is computed using the same methodology as the 30-day VIX index.

[18]The path displayed in Figure 2 differs from the one examined in previous studies (e.g., Bekaert and Hoerova (2014), Carr and Wu (2009)) and computed as $\hat{E}(rv_{t+1}| I_t) - iv_t$, where $\hat{E}(rv_{t+1}| I_t)$ is an estimate of the conditional expected realized variance and iv_t is the implied variance. Here, we project

purposes, we also plot the equity VRP previously depicted in Figure 1. The option VRP projection is generally negative and exhibits pronounced spikes. For instance, the magnitude of the premium reaches 4.96% per year in 1998 (LTCM collapse and Russian crisis), 11.07% in 2008 (height of the financial crisis), and 6.40% in 2011 (European debt crisis).

[TABLE 5 HERE]

[FIGURE 2 HERE]

4.2 The Variance Risk Premium Difference

The two premia in Figure 2 share strong similarities and have a correlation coefficient of 0.67. However, they also exhibit large discrepancies which take both positive and negative values. For instance, the magnitude of the option VRP is substantially larger during the 2008 and European debt crises, whereas a negative VRP difference is observed during the late 1990s and the early 2000s. In addition, the VRP difference is persistent over time (i.e., its autocorrelation coefficient equals 0.48), which implies that several quarters are necessary for the pricing gap between the two markets to vanish.

Given our previous analysis, we expect this difference to be mostly driven by the broker-dealer variables (leverage and PBI return). To formally address this issue, we compute the estimated vector $\hat{V}_v^o - \hat{V}_v^e$ that drive the time-variation of the VRP difference. The results in Panel B of Table 6 confirm that the two broker-dealer variables play a central role. Regarding the leverage ratio, the estimated coefficient is highly significant and implies that a one-standard deviation decline in leverage increases the gap between the equity and option option VRPs by -1.80% per year ($-0.45 \cdot 4$)—a change nearly as large as the average premium itself. A similar result holds for the PBI return which yields a negative and significant coefficient of -0.19. A simple illustration of these results is provided by Figure 3 which plots the VRP difference (black line), alongside with the

both $E\left(rv_{t+1}|\, I_t\right)$ and iv_t on the linear space spanned by z_t to allow for a meaningful comparison between the equity and option markets.

quarterly leverage ratio of intermediaries (dashed line). We see that periods when leverage increases (decreases) coincide with a decline in the magnitude of the option VRP projection relative to that of the equity market.

Panel A reveals that the macro-finance variables are not relevant for explaining the VRP difference. In addition, we observe that the average difference is close to zero (-0.09% per quarter). Therefore, a simple analysis of the unconditional premia is insufficient to uncover the large, but temporary discrepancies between the two markets.

In summary, the empirical evidence documented here reveals that the equity and option VRP projections are, on average, identical and respond similarly to changes in economic and volatility conditions. However, their sensitivities to the broker-dealer variables differ dramatically: the leverage ratio and PBI return are strongly related to the option VRP, but leave the equity VRP nearly unchanged. Therefore, both predictors signal periods when the price of variance risk differs across the two markets.

[TABLE 6 HERE]

[FIGURE 3 HERE]

4.3 Equity-Based Misspecification Tests

4.3.1 Specification Test of the Two-Factor Model

As discussed in Section 2, the estimation procedure requires the two-factor model to capture the return dynamics of the 25 equity portfolios.[19] If it is not the case, the equity vector $\hat{V}_v^e$ can be biased, resulting in a difference between the equity and option markets. To address this issue, the most direct approach is to formally test the two-factor model. Under the null hypothesis of correct specification, we know from equation (8) that c_p is equal to $b_{pv} \cdot V_v^e + b_{pm} \cdot V_m^e$. Therefore, we can perform a joint test across the 25 portfolios based on the sum of the squared pricing errors, $Q = \sum_{p=1}^{25} (c_p - B_p V^e)' (c_p - B_p V^e)$, where B_p a $J \times 2J$ matrix equal to $[b_{pv} \cdot I_J, b_{pm} \cdot I_J]$, I_J is a $J \times J$ identity matrix,

[19]Note that the two-factor is only required to correctly price the 25 equity portfolios, not the entire cross-section of individual stocks.

and $V^e = [V_v^{e\prime}, V_m^{e\prime}]'$.[20] Table 4 reveals that the test statistic (J-stat) is far below the conventional rejection thresholds with or without the broker-dealer variables (the p-values range between 0.32 and 0.43).

A second and related specification check is to compare, for each portfolio, (i) the unconstrained projection, $proj(r_{p,t+1}^e \mid z_t)$, from a linear regression of $r_{p,t+1}^e$ on z_t with (ii) the version constrained by the model and defined as $proj^M(r_{p,t+1}^e \mid z_t) = b_{pv}\lambda_{v,t}^e(z) + b_{pm}\lambda_{m,t}^e(z)$, where $\lambda_{m,t}^e(z)$ is the market risk premium projection. If the model is correctly specified, we expect a close match between the two projections. The results in the appendix reveal that the R^2 from a time-series regression of the estimated value of $proj(r_{p,t+1}^e \mid z_t)$ on that of $proj^M(r_{p,t+1}^e \mid z_t)$ reaches 98.2% on average, which provides further evidence in favor of the two-factor model.

The third test consists in studying the estimated market premium $\widehat{\lambda}_{m,t}^e(z) = (\hat{F}_m - \hat{V}_m^e)'z_t$, where $\hat{F}_m$ is the coefficient vector from the predictive regression of the market factor and $\hat{V}_m^e$ is inferred from the conditional two-pass regression. If the model is correctly specified, the vector V_m^e must be equal to zero because the market factor is an excess return which has, by construction, a zero-price (i.e., $\lambda_{m,t}^e(z) = F_m'z_t$). The results reported in the appendix confirm that no element of the estimated vector $\hat{V}_m^e$ is significantly different from zero at conventional levels. In addition, the estimated premium exhibits the traditional properties documented in the previous literature as it is countercyclical and strongly related to the PE ratio (e.g., Fama and French (1989) and Keim and Stambaugh (1986)).

Finally, we examine the linearity assumption by measuring the ability of the 25 equity portfolios to track the realized variance payoff. We construct a mimicking portfolio that maximizes the conditional correlation with rv_{t+1} by applying the method developed by Ferson, Siegel, and Xu (2006) and Lamont (2001). The portfolio payoff is defined as $b'r_{t+1}^e$, where r_{t+1}^e is the vector of portfolio returns, and b is the coefficient vector from the following time-series regression: $rv_{t+1} = c'z_t + b'r_{t+1}^e + e_{t+1}$, where z_t includes the macro-

[20] The formal test statistic and its asymptotic distribution derived by Kan, Robotti, and Shanken (2013) are described in the appendix.

finance predictors. For comparison purposes, we repeat this procedure with the excess return of the replicating option strategy r^o_{t+1} computed using the approach of Carr and Wu (2009). The correlation produced by the equity-based portfolio reaches 0.82, and is nearly as high as the level observed in the option market (0.93). To visualize these results, we display in Figure 4 the close relationships between the innovation of the realized variance and the innovations of the equity- and option-based mimicking portfolios.

[FIGURE 4 HERE]

4.3.2 The Realized Variance Components

The two-factor model includes the realized variance but it does not distinguish between its different components. For instance, Todorov (2010) distinguishes between the diffusion and jump components of the realized variance, while Adrian and Rosenberg (2008) measure its low- and high-frequency components. Because the 25 equity portfolios may load very differently on each component, one concern is that bundling them into a single factor may potentially bias the equity vector $\hat{V}^e_v$. To demonstrate that it is not the case, let us focus on the coefficient for the leverage ratio, $v^e_v(lev)$, in a three-factor model that includes the market factor $f_{m,t+1}$ and the two components of the realized variance, denoted by $rv_{1,t+1}$ and $rv_{2,t+1}$. This model can be conveniently written as $r^{e,*}_{p,t+1} = c_{p,0} + c_{p,lev} \cdot lev^*_{t-1} + b_{pv,1} \cdot rv^*_{1,t+1} + b_{pv,2} \cdot rv^*_{2,t+1} + e_{p,t+1}$, where the star notation denotes the residual from the projection on the space spanned by $f_{m,t+1}$ and z_t (except leverage). Using the equilibrium restriction $c_{p,lev} = b_{pv,1} \cdot v^e_{v,1}(lev) + b_{pv,2} \cdot v^e_{v,2}(lev)$, we can estimate $v^e_{v,1}(lev)$, $v^e_{v,2}(lev)$ and add both terms to obtain an unbiased estimator of $v^e_v(lev)$. Alternatively, we can replace $rv_{1,t+1}$ and $rv_{2,t+1}$ with its sum and rewrite the model as $r^{e,*}_{p,t+1} = c_{p,0} + c_{p,lev} \cdot lev^*_{t-1} + b_{pv} \cdot rv^*_{t+1} + e_{p,t+1}$, where b_{pv} is a weighted average of $b_{pv,1}$ and $b_{pv,2}$.[21] The equilibrium condition becomes $c_{p,lev} = b_{pv}(v^e_{v,1}(lev) + v^e_{v,2}(lev)) = b_{pv} \cdot v^e_v(lev)$,

[21] Specifically, we can write b_{pv} as $\frac{cov(r^{e,*}_{p,t+1}, rv^*_{t+1} | lev^*_t)}{var(rv^*_{t+1} | lev^*_t)} = \frac{(w_1 b_{pv,1} + w_2 b_{pv,2})}{var(rv^*_{t+1} | lev^*_t)}$, where $w_1 = \frac{var(rv^*_{1,t+1} | lev^*_t, rv^*_{2,t+1}) cov(r^{e,*}_{p,t+1}, rv^*_{1,t+1} | lev^*_t)}{cov(r^{e,*}_{p,t+1}, rv^*_{1,t+1} | lev^*_t, rv^*_{2,t+1})}$, and $w_2 = \frac{var(rv^*_{2,t+1} | lev^*_t, rv^*_{1,t+1}) cov(r^{e,*}_{p,t+1}, rv^*_{2,t+1} | lev^*_t)}{cov(r^{e,*}_{p,t+1}, rv^*_{2,t+1} | lev^*_t, rv^*_{1,t+1})}$.

which again allows for an unbiased estimation of $v_v^e(lev)$.

To validate this analysis empirically, we build on Adrian and Rosenberg (2008) and re-estimate the vector $F_v - V_v^e$ using a three-factor model that includes the low- and high-frequency variance components. The coefficients reported in the appendix are essentially the same as those in Table 4.

4.3.3 Time-Varying Portfolio Betas

Next, we formally examine the potential impact of time-varying betas on the equity vector $\hat{V}_v^e$. To begin, we allow the variance and market betas to depend linearly on each predictor $z_{j,t-1}$: $b_{pv,t} = b_{pv,0} + b_{pv,j} z_{j,t-1}$ and $b_{pm,t} = b_{pm,0} + b_{pm,j} z_{j,t-1}$ (for $j = 1, ..., J$).[22] Under this specification, the model derived in the appendix contains two additional factors, $f_{1,t+1} = z_{j,t-1} rv_{t+1}$ and $f_{2,t+1} = z_{j,t-1} f_{m,t+1}$, whose omission can possibly bias the vector $\hat{V}_v^e$. In short, the appendix reveals that only 13% of the 350 estimated coefficients of $b_{pv,j}$ and $b_{mv,j}$ are statistically significant (at the 5% level). Therefore, we find little empirical evidence to support the view that betas change with the predictive variables.

We also consider the case where portfolio betas are subject to a structural break. To elaborate, suppose that the variance beta of each portfolio is equal to $b_{pv,1}$ for $t \leq \tau$ and then moves to $b_{pv,2}$. Applying the same demonstration as the one presented above for the variance components reveals that the full-sample estimation of the equity vector V_v^e remains valid.[23] However, conducting a subperiod analysis is useful for several reasons. First, we can verify that the monotonicity of the variance betas is a robust feature of the equity portfolios. This is confirmed in Table 7 which reports the variance betas over the periods 1970-1991 and 1992-2012. Second, we can test whether the difference between the equity and option markets is not an artefact of our econometric treatment of samples of unequal lengths. In Table 8, we report the vector $\hat{V}_v^o - \hat{V}_v^e$ that drive the VRP difference

[22] A full-fledged model with all predictors is subject to the curse of dimensionality because more than 50 coefficients must be estimated for each portfolio (see Gagliardini, Ossola, and Scaillet (2014)).

[23] We simply need to define $rv_{1,t+1}$ as $rv_{t+1} \cdot 1\{t \leq \tau\}$ and $rv_{2,t+1}$ as $rv_{t+1} \cdot 1\{t > \tau\}$, where $1\{\cdot\}$ is an indicator function.

using data from the short sample only (1992-2012). Consistent with our initial analysis, the relationships between the broker-dealer variables (leverage, change in leverage, PBI return) remain negative and statistically significant.

[TABLE 7 HERE]

[TABLE 8 HERE]

4.3.4 Additional Risk Factors

Finally, we examine whether the VRP difference could be caused by the omission of relevant equity risk factors. We re-estimate the equity vector V_v^e using extended versions of the two-factor model that include: (i) size and Book-to-Market (BM) factors; (ii) size, BM, and momentum factors; (iii) size, BM, and liquidity factors; (iv) the squared realized variance to allow for a non-linear (quadratic) relationship between the portfolio returns and the realized variance. The results reported in the appendix reveal that the estimated coefficients are all similar to those presented here.

In summary, we find that formal specification tests do not reject the two-factor model. We also demonstrate that grouping the different variance components into a single factor do not bias the estimated coefficients. Finally, modeling time-varying betas and including additional sources of risk leave the results nearly unchanged. The empirical evidence therefore suggests that equity-based misspecification is not the driving force behind the observed difference between the equity and option markets.

5 Interpreting the Evidence

5.1 Discrepancies between the Equity and Option Markets

The main insight from the empirical analysis is that the VRPs in the equity and option markets can be significantly different, both statistically and economically. The implications of this result are twofold. First, it leads to a more nuanced view of the information content of the option VRP. Because the latter can be easily computed from the VIX index, it is a widely-used proxy for fluctuations in the risk aversion of equity investors. Despite the commonalities between the two markets displayed in Figure 2, this interpretation can be misleading because the option VRP is disproportionately influenced by the broker-dealer variables. Spikes in the option VRP can arise when broker-dealers are in a deleveraging phase, while a low price of variance risk could be the consequence of their increased willingness to take on risk. Yet, these variations do not necessarily mean that investors in the equity market change their attitude towards stocks exposed to variance risk.

To further corroborate this assertion, we perform a simple test in which we examine the predictive ability of the broker-dealer variables on the market return. If these variables contain information that is not directly related to the risk aversion of equity investors, their forecasting ability should be weak in the presence of the macro-finance variables. We find that none of the relationships is statistically significant and that the coefficient for the PBI return has the wrong sign (i.e., the t-statistics for leverage and PBI return are equal to -0.99 and 0.74, respectively).

Second, the rejection of the null hypothesis of equal VRPs indicates the presence of some market frictions that prevent the law of one price to apply. Such constraints generate a shadow cost on investors' utility which results in different prices of variance risk in equilibrium. This constrained environment therefore departs from the model proposed by Chen, Joslin, and Ni (2013) in which shocks to the risk-bearing capacity of financial intermediaries affect option prices and are also fully reflected in the equity market.

5.2 Possible Explanations Based on Market Frictions

The international finance literature commonly interprets mispricing across countries as evidence of market segmentation (e.g., Bekaert, Harvey, Lundblad, and Siegel (2011)).

Consistent with this view, the VRP difference can possibly be caused by informational or regulatory constraints that limit risk-sharing between marginal investors in the equity and option markets. The theoretical motivation is provided by Basak and Croitoru (2000) who demonstrate how deviations from the law of one price exist in equilibrium in the presence of portfolio constraints that limit investors' positions in the two markets. In practice, these constraints can take several forms. Retail investors may lack the expertise required to monitor option positions, whereas mutual funds generally face limits on the amount of options they can hold in their portfolios. In addition, option trading desks generally have the mandate to trade exclusively in the underlying necessary to manage the delta of their option positions (i.e., in index futures), but not in stocks exposed to the variance factor. Under this scenario, when risk-constrained intermediaries deleverage and the option VRP is high (in absolute value), equity investors are unable to write options in sufficient number to provide protection against spikes in aggregate volatility.[24] Conversely, when the option VRP is low (in absolute value), stock market investors do not fully exploit low option prices and broker-dealers fail to aggressively trade in stocks to reduce the magnitude of the equity VRP.

Alternatively, the gap between the two markets could be driven by margin requirements. Garleanu and Pedersen (2011) demonstrate that identical assets can exhibit different prices if they are traded in markets in which margins differ. Applied to our setting, their theory predicts that the price of identical cash flows should be lower in the stock market because it exhibits higher margin requirements than the option market. Furthermore, this price discrepancy should increase in the tightness of funding constraints, leading to a time-varying and positive VRP difference between the equity and option markets.

[24] Anecdotal evidence suggests that very few equity investors wrote put options during the recent crisis, despite the fact that they were highly priced. One notable exception is Warren Buffet whose short positions in equity put options reached a notional size of $35-40 billion in 2008 (Triana (2013)). A key reason for building this option position is that Buffet secured a deal in which puts were not marked-to-market in case of adverse market movements. Therefore, Buffet benefitted from a special treatment that is not available to most investors.

Whereas both explanations based on segmentation and margin requirements are likely to play a role, the second cannot be fully reconciled with the path followed by the VRP difference for two reasons. First, it cannot easily account for the positive and negative VRP differences observed in Figure 3 because margin requirements in the option market are unlikely to be greater than those in the equity market. Second, under the margin-based story, the explanatory power of the broker-dealer variables stems from their ability to track changes in funding constraints. However, Table 9 shows that alternative and, arguably, more direct measures of funding constraints are not strongly related to the VRP difference. The coefficient associated with the default spread has the right sign but is not statistically significant (in both univariate and multivariate regressions). On the contrary, the TED spread produces a coefficient with the wrong sign, i.e., the VRP difference drops when funding constraints become tighter.

[TABLE 9 HERE]

5.3 Broker-Dealer Variables and Option Supply

The strong explanatory power of the broker-dealer variables on the option VRP resonates with the key role played by these intermediaries in the option market. Chen, Joslin, and Ni (2013), Fournier (2014), and Garleanu, Pedersen, and Poteshman (2009) empirically demonstrate that public investors have a long net position in SP500 index options, particularly in deep out-of-the-money put options. By market clearing, financial intermediaries write options to satisfy this demand and are structurally short variance risk. As a result, these authors argue that option prices are determined by local supply and demand factors. In particular, changes in intermediaries' risk-bearing capacity should move the option supply curve and affect option prices.

To test the validity of this supply-based mechanism, we can examine the relationships between the broker-dealer variables and option prices. Provided that high leverage and PBI return signal a high risk-bearing capacity (Adrian and Shin (2010, 2013)), both variables should have a negative impact on option prices. In Table 10, we report the estimated vector $\hat{V}_v^o$ from the implied variance regression. Because the implied variance is a measure of option expensiveness, $\hat{V}_v^o$ can be interpreted as the option price reaction to changes in the predictor values. The results in Panel B provide evidence in favor of supply effects, i.e., the coefficients are all strongly negative (-0.07 and -0.14) and imply that options become cheaper when the leverage ratio and PBI return are high.[25]

There are two potential concerns with the previous interpretation. First, the leverage ratio may also measure the quantity of options exchanged in the market. In this case, it should be treated as an endogenous variable determined along with the option price. In a endogenous price-quantity regression, Hamilton (1994) demonstrates that the slope coefficient (i) is a mixture of the negative demand slope and the positive supply slope, and (ii) is negative when supply shocks are the main determinants of the traded price and quantity. Therefore, the negative coefficient in Table 10 still provides supporting evidence of a supply-based mechanism. Second, the coefficients associated with the broker-dealer variables could affected by the omission of a relevant variable. While this case cannot be definitively ruled out, the set of predictors examined in Table 10 includes a range of macro-finance variables. As discussed below, we also consider additional predictors which all leave the explanatory power of the broker-dealer variables unchanged.[26]

[TABLE 10 HERE]

[25] Interestingly, we also observe that the leverage of broker-dealers increases when the target federal funds rate drops, which implies that monetary easing is associated with greater risk taking by financial intermediaries, and with a lower VRP in the option market. This finding resonates with the model recently proposed by Drechsler, Savov, and Schnabl (2014), in which lower nominal rates result in increased bank leverage and lower risk premia.

[26] Cheng, Kirilenko, and Xiong (2012) and Etula (2013) also provide empirical evidence that the risk-bearing capacity of financial intermediaries has an impact on prices in commodity futures and derivatives markets. An important difference with these studies is that we control for a large set of macro-finance variables that could potentially affect option prices.

6 Additional Results

6.1 Alternative Set of Predictive Variables

We re-estimate the coefficients associated with the two broker-dealer variables in the presence of additional macro-finance variables, namely the dividend yield, the quarterly growth rate in industrial production, the business cycle indicator constructed by Aruoba, Diebold, and Scotti (2009), the 3-month T-bill rate, the term spread, and the quarterly volatility of inflation. In addition, we examine whether the broker-dealer variables capture non-linear effects by including the squared value of each macro-finance variable. In all of these cases, the estimated coefficients for the leverage and PBI return are statistically insignificant in the equity market, but strongly significant in the option market (see the appendix).

6.2 Extreme Variance Observations

The differential impact of the two broker-dealer variables on the equity and option markets could possibly be driven by a few outlier observations of the variance process. To address this issue, we winsorize 2% and 5% of the most extreme market variance datapoints (1% and 2.5% in each tail) and re-estimate each VRP projection. In both cases, we still document a strong and statistically significant difference between the two markets (see the appendix).

6.3 Alternative Measure of the Implied Variance

The procedure for estimating the option VRP requires as input the implied variance measured as the squared VIX index computed from option prices. However, the equality between the two variables does not hold perfectly when the market is subject to large movements. To address this issue, we build on recent work by Martin (2013) and consider the SVIX index that is robust to jumps.[27] Consistent with the results in Table 10, the appendix reveals that the broker-dealer variables remain negatively correlated with the SVIX index.

7 Conclusion

In this paper, we formally compare the conditional VRPs inferred from equity and option prices. We find that the premia in both markets are, on average, in line with one another and respond similarly to changes in volatility and business cycle conditions. However, we identify episodes when they diverge and find that such differences are explained to a large extent by two broker-dealer variables that measure the financial standing of intermediaries. An increase (decrease) in the leverage or past performance of intermediaries decreases (increases) the magnitude of the option VRP, but leaves the equity VRP unchanged.

The rejection of the null hypothesis that the two VRPs are equal implies that caution should be exercised when the option VRP is used as a measure of risk aversion of equity investors. In addition, it also indicates the presence of frictions between the two markets that prevent the law of one price to apply. Finally, the close relationships between the broker-dealer variables and the option VRP is consistent with the key role played by intermediaries in the option market.

These results can be exploited in future theoretical work that attempts to explain the aggregate pricing of variance risk and model local demand and supply factors in the option market. They also provide novel empirical evidence regarding the connection between risk-taking by financial intermediaries and asset prices. Understanding the nature of this connection is a major concern for policymakers (e.g., Bernanke and Kuttner (2005), Rajan (2006)) and an interesting avenue of future research.

[27] We thank Ian Martin for providing us with the data.

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Table 1: Summary Statistics for the Predictive Variables

Panel A reports the quarterly mean and standard deviation of the different variables used to explain the dynamics of the variance risk premium, which are the lagged realized variance, the price/earnings ratio, the default spread, the quarterly inflation rate of the producer price index, the quarterly growth rate of the number of employees in the nonfarm sector (seasonally-adjusted), the leverage ratio of broker-dealers, and the quarterly return of the prime broker index (all expressed in log form). The remaining columns of Panel A show the skewness, kurtosis, first-, and second-order partial autocorrelation coefficients of the standardized versions of the predictors. Panel B shows the correlation matrix of the standardized predictors. All statistics are computed using quarterly data between January 1970 and September 2012 (171 observations).

Panel A: Unconditional Moments

	Mean	Std.	Skew.	Kurt.	AC1	AC2
Lagged Realized Variance (RV)	−5.32	0.80	0.82	4.43	0.65	0.13
Price/Earnings Ratio (PE)	1.24	0.20	0.14	2.18	0.93	−0.10
Default Spread (DEF)	1.03%	0.41%	2.16	10.81	0.81	−0.11
Producer Price Index (PPI)	0.94%	1.31%	0.00	5.69	0.34	0.18
Employment Growth (EMP)	0.37%	0.58%	−0.99	4.74	0.75	0.03
Broker-Dealer Leverage (LEV)	2.70	0.61	0.84	4.50	0.85	0.05
Prime Broker Index (PBI)	1.78%	17.9%	−0.54	4.45	0.05	−0.14

Panel B: Correlations

	PE	DEF	PPI	EMP	LEV	PBI
Lagged Realized Variance (RV)	−0.03	0.46	−0.06	−0.42	0.05	−0.29
Price/Earnings Ratio (PE)		−0.52	−0.11	0.23	0.33	0.07
Default Spread (DEF)			−0.21	−0.62	0.05	−0.01
Producer Price Index (PPI)				0.09	−0.11	−0.02
Employment Growth (EMP)					−0.12	−0.05
Broker-Dealer Leverage (LEV)						−0.14

Table 2: Summary Statistics for the Variance Portfolios

Panel A shows the annualized excess mean, standard deviation, size (in log form), book-to-market ratio, and the pre-, post-rank variance betas of the quarterly returns of quintile portfolios formed by equally weighting all variance portfolios in the same variance beta quintile (Low, 2, 3, 4, High). For each portfolio, the pre-rank beta is defined as the mean of the variance beta across stocks on the portfolio formation date over the whole sample. The post-rank variance beta is computed from the time-series regression of the portfolio return on the realized variance, the market return, and the macro-finance variables. Panel B reports the annualized estimated alpha of each quintile portfolio using the CAPM, the Fama-French (FF) model based on market, size, and book-to-market factors, and two FF-extensions that include momentum and liquidity portfolios, respectively. The t-statistics of the different estimators are shown in parentheses and are robust to the presence of heteroskedasticity. ***, **, and * designate statistical significance at the 1%, 5%, and 10% levels, respectively.

Panel A: Unconditional Moments, Characteristics, and Variance Betas

Quintile	Mean (% p.a.)	St. Dev. (% p.a.)	Size	BM	Pre-rank beta		Post-rank beta	
Low	7.42	17.76	8.14	0.73	−0.64	(−2.58)	−0.61***	(−2.99)
2	7.00	17.44	8.41	0.69	−0.28	(−0.88)	−0.45***	(−2.88)
3	6.57	17.21	8.60	0.67	0.02	(0.02)	−0.15	(−0.78)
4	5.46	17.86	8.63	0.67	0.31	(0.84)	−0.09	(−0.27)
High	4.44	17.69	8.48	0.68	0.67	(2.58)	0.16	(0.50)
High-Low	−2.99	8.02	0.33	−0.05	1.31	(5.16)	0.76**	(2.50)

Panel B: Alphas

Quintile	CAPM (% p.a.)		Fama-French (FF) (% p.a.)		FF+Mom (% p.a.)		FF+Liq (% p.a.)	
Low	2.00*	(1.91)	0.80	(0.74)	0.80	(0.74)	0.40	(0.39)
2	1.40*	(1.84)	0.25	(0.38)	0.08	(0.11)	0.22	(0.32)
3	1.04	(1.35)	0.11	(0.16)	0.10	(0.15)	0.07	(0.10)
4	−0.32	(−0.45)	−1.21*	(−1.67)	−1.23*	(−1.65)	−1.28*	(−1.80)
High	−1.20	(−1.38)	−2.40***	(−2.69)	−2.00*	(−1.90)	−2.00**	(−2.18)
High-Low	−3.20***	(−2.51)	−3.20**	(−2.37)	−2.80*	(−1.89)	−2.40*	(−1.87)

Table 3: Realized Variance Predictability

Panel A reports the estimated coefficients and the predictive R^2 of a time-series regression of the quarterly realized variance on the set of macro-finance predictors that includes the lagged realized variance (RV), the price/earnings ratio (PE), the default spread (DEF), the quarterly inflation rate (PPI), and the quarterly employment rate (EMP). The coefficients determine the impact of a one-standard deviation change in the predictors on the future realized variance. The figures in parentheses report the t-statistics of the estimated coefficients that are robust to the presence of heteroskedasticity. Panel B examines the incremental predictive power of the broker-dealer variables. The leftmost columns show the results for the broker-dealer leverage ratio (LEV) and the quarterly return of the prime broker index (PBI) in the presence of the macro-finance variables. The rightmost columns repeat the analysis after replacing the leverage ratio with the annual change in the leverage ratio (ΔLEV). ***, **, and * designate statistical significance at the 1%, 5%, and 10% levels, respectively.

Panel A: Macro-Finance Variables

	Mean	R. Var. (RV)	PE Ratio (PE)	Default (DEF)	Inflation (PPI)	Employ. (EMP)	R^2
R. Variance	0.75***	0.48***					0.16
	(8.78)	(3.49)					
All Variables	0.75***	0.39***	0.18*	0.25**	0.12	0.02	0.18
	(8.89)	(3.52)	(1.84)	(2.21)	(1.12)	(0.32)	

Panel B: Contribution of Broker-Dealer Variables

	Leverage (LEV)	PB Index (PBI)	R^2	ΔLeverage (ΔLEV)	PB Index (PBI)	R^2
+Leverage	0.28		0.23	0.19		0.20
	(1.19)			(0.99)		
+Prime Broker		−0.07	0.18		-	-
		(−1.10)			-	
+Leverage & Prime Broker	0.28	0.01	0.23	0.19	−0.05	0.21
	(1.08)	(0.05)		(0.93)	(−0.61)	

Table 4: Equity Variance Risk Premium

Panel A reports the estimated coefficients that drive the equity Variance Risk Premium (VRP) for the set of macro-finance predictors that includes the lagged realized variance (RV), the price/earnings ratio (PE), the default spread (DEF), the quarterly inflation rate (PPI), and the quarterly employment rate (EMP). The coefficients determine the impact of a one-standard deviation change in the predictors on the VRP, and are computed using the conditional two-pass regression described in Section 2. The figures in parentheses report the t-statistics of the estimated coefficients that are robust to the presence of heteroskedasticity. The J-statistic and associated p-values in brackets are based on the joint test proposed by Kan, Robotti, and Shanken (2013) and described in the appendix. Panel B examines the incremental predictive power of the (standardized) broker-dealer variables. The leftmost columns show the results for the broker-dealer leverage ratio (LEV) and the quarterly return of the prime broker index (PBI) in the presence of the macro-finance variables. The rightmost columns repeat the analysis after replacing the leverage ratio with the annual change in the leverage ratio (ΔLEV). ***, **, and * designate statistical significance at the 1%, 5%, and 10% levels, respectively.

Panel A: Macro-Finance Variables

	Mean	R. Var. (RV)	PE ratio (PE)	Default (DEF)	Inflation (PPI)	Employ. (EMP)	J-stat.
R. Variance	-0.58^{***}	-0.33^{*}					
	(−2.78)	(−1.76)					
All Variables	-0.58^{***}	-0.34^{*}	0.16	0.20	0.32^{*}	0.18	4.88
	(−2.84)	(−1.81)	(0.85)	(0.77)	(1.76)	(0.87)	[0.41]

Panel B: Contribution of Broker-Dealer Variables

	Leverage (LEV)	PB Index (PBI)	J-stat.	ΔLeverage (ΔLEV)	PB Index (PBI)	J-stat.
+Leverage	0.01		5.88	−0.03		6.05
	(0.05)		[0.34]	(−0.21)		[0.31]
+Prime Broker		−0.02	5.70		-	-
		(−0.11)	[0.43]		-	-
+Leverage & Prime Broker	−0.04	−0.02	6.85	−0.03	−0.02	6.86
	(−0.21)	(−0.11)	[0.32]	(−0.18)	(−0.15)	[0.33]

Table 5: Option Variance Risk Premium

Panel A reports the estimated coefficients that drive the option Variance Risk Premium (VRP) for the set of macro-finance predictors that includes the lagged realized variance (RV), the price/earnings ratio (PE), the default spread (DEF), the quarterly inflation rate (PPI), and the quarterly employment rate (EMP). The coefficients determine the impact of a one-standard deviation change in the predictors on the VRP, and are computed using the GMM approach described in Section 2. The figures in parentheses report the t-statistics of the estimated coefficients that are robust to the presence of heteroskedasticity. Panel B examines the incremental predictive power of the broker-dealer variables. The leftmost columns show the results for the broker-dealer leverage ratio (LEV) and the quarterly return of the prime broker index (PBI) in the presence of the macro-finance variables. The rightmost columns repeat the analysis after replacing the leverage ratio with the annual change in the leverage ratio (ΔLEV). ***, **, and * designate statistical significance at the 1%, 5%, and 10% levels, respectively.

Panel A: Macro-Finance Variables

	Mean	R. Var. (RV)	PE Ratio (PE)	Default (DEF)	Inflation (PPI)	Employ. (EMP)
R. Variance	-0.47^{***}	-0.31^{***}				
	(-7.04)	(-3.00)				
All Variables	-0.47^{***}	-0.37^{***}	0.30^{***}	0.03	0.17^{**}	-0.09
	(-7.66)	(-3.73)	(4.11)	(0.22)	(2.07)	(-1.03)

Panel B: Contribution of Broker-Dealer Variables

	Leverage (LEV)	PB Index (PBI)	ΔLeverage (ΔLEV)	PB Index (PBI)
+Leverage	0.35^{***}		0.27^{***}	
	(3.93)		(3.73)	
+Prime Broker		0.07		-
		(0.80)		-
+Leverage & Prime Broker	0.42^{***}	0.18^{**}	0.34^{***}	0.14^{*}
	(4.65)	(1.96)	(5.02)	(1.79)

Table 6: Equity versus Option Variance Risk Premia

Panel A reports the estimated coefficients that drive the difference between the equity and option Variance Risk Premia (VRPs) for the set of macro-finance variables that includes the lagged realized variance (RV), the price/earnings ratio (PE), the default spread (DEF), the quarterly inflation rate (PPI), and the quarterly employment rate (EMP). The coefficients determine the impact of a one-standard deviation change in the predictors on the difference between the equity and option VRPs. The figures in parentheses report the t-statistics of the estimated coefficients computed using a bootstrap procedure described in the appendix. Panel B examines the incremental predictive power of the broker-dealer variables. The leftmost columns show the results for the broker-dealer leverage ratio (LEV) and the quarterly return of the prime broker index (PBI) in the presence of the macro-finance variables. The rightmost columns repeat the analysis after replacing the leverage ratio with the annual change in the leverage ratio (ΔLEV). ***, **, and * designate statistical significance (based on the bootstrap distributions) at the 1%, 5%, and 10% levels, respectively.

Panel A: Macro-Finance Variables

	Mean	R. Var. (RV)	PE ratio (PE)	Default (DEF)	Inflation (PPI)	Employ. (EMP)
R. Variance	−0.09	−0.02				
	(−0.50)	(−0.11)				
All Variables	−0.09	0.03	−0.14	0.18	0.15	0.26
	(−0.34)	(0.13)	(−0.82)	(0.77)	(1.02)	(1.52)

Panel B: Contribution of Broker-Dealer Variables

	Leverage (LEV)	PB Index (PBI)	ΔLeverage (ΔLEV)	PB Index (PBI)
+Leverage	-0.34^{***}		-0.30^{**}	
	(−3.17)		(−2.46)	
+Prime Broker		−0.09		-
		(−1.05)		-
+Leverage & Prime Broker	-0.45^{***}	-0.19^{**}	-0.37^{***}	-0.17^{*}
	(−4.19)	(−2.07)	(−3.06)	(−1.68)

Table 7: Variance Betas: Subperiod Analysis

This table reports the pre-, post-rank variance betas of the quarterly returns of quintile portfolios formed by equally weighting all variance portfolios in the same variance beta quintile (Low, 2, 3, 4, High) over the periods 1970-1991 and 1992-2012. For each portfolio, the pre-rank beta is defined as the mean of the variance beta across stocks on the portfolio formation date over the whole sample. The post-rank variance beta is computed from the time-series regression of the portfolio return on the realized variance, the market return, and the macro-finance predictors. The t-statistics of the different estimators are shown in parentheses and are robust to the presence of heteroskedasticity. *** and ** designate statistical significance at the 1% and 5% levels, respectively.

	Period 1970-1991				Period 1991-2012			
Quintile	Pre-rank beta		Post-rank beta		Pre-rank beta		Post-rank beta	
Low	−0.68	(−2.73)	−0.09	(−0.59)	−0.59	(−2.41)	-0.60^{***}	(−3.00)
2	−0.32	(−0.79)	−0.02	(−0.35)	−0.23	(−0.79)	-0.45^{***}	(−2.88)
3	0.00	(0.02)	0.34^{***}	(4.28)	0.03	(0.02)	−0.15	(−0.77)
4	0.33	(0.85)	0.73^{***}	(5.76)	0.29	(0.85)	−0.08	(−0.27)
High	0.73	(2.50)	0.90^{***}	(7.78)	0.61	(2.50)	0.16	(0.50)
High-Low	1.42	(5.39)	0.98^{***}	(5.56)	1.20	(4.92)	0.77^{**}	(2.50)

Table 8: Variance Risk Premium Difference: Short Sample

Panel A reports the estimated coefficients that drive the equity Variance Risk Premium (VRP) over the short sample (1992-2012) for the set of macro-finance predictors that includes the lagged realized variance (RV), the price/earnings ratio (PE), the default spread (DEF), the quarterly inflation rate (PPI), and the quarterly employment rate (EMP). The coefficients determine the impact of a one-standard deviation change in the predictors on the VRP, and are computed using the conditional two-pass regression described in Section 2. The figures in parentheses report the t-statistics of the estimated coefficients that are robust to the presence of heteroskedasticity. Panel B examines the incremental predictive power of the broker-dealer variables. The leftmost columns show the results for the broker-dealer leverage ratio (LEV) and the quarterly return of the prime broker index (PBI) in the presence of the macro-finance variables. The rightmost columns repeat the analysis after replacing the leverage ratio with the annual change in the leverage ratio (ΔLEV). ***, **, and * designate statistical significance at the 1%, 5%, and 10% level, respectively.

Panel A: Macro-Finance Variables

	Mean	R. Var. (RV)	PE Ratio (PE)	Default (DEF)	Inflation (PPI)	Employ. (EMP)
R. Variance	−0.58	−0.15				
	(−1.29)	(−0.72)				
All Variables	−0.58	0.17	−0.41*	−0.27	0.18	0.27
	(−1.48)	(0.70)	(1.76)	(0.83)	(0.86)	(1.12)

Panel B: Contribution of Broker-Dealer Variables

	Leverage (LEV)	PB Index (PBI)	ΔLeverage (ΔLEV)	PB Index (PBI)
+Leverage	−0.55***		−0.38**	
	(−3.88)		(2.41)	
+Prime Broker		−0.05		—
		(−0.41)		—
+Leverage & Prime Broker	−0.59***	−0.18*	−0.42**	−0.15
	(−4.09)	(−1.66)	(2.44)	(1.18)

Table 9: Variance Risk Premium Difference: Proxies for Funding Constraints

This table reports the estimated coefficients that drive the difference between the equity and option Variance Risk Premia (VRPs) for two proxies of funding constraints, namely the default spread (DEF) and the TED spread (TED). The explanatory power of these variables is measured with and without the set of macro-finance predictors that includes the lagged realized variance (RV), the price/earnings ratio (PE), the quarterly inflation rate (PPI), and the quarterly employment rate (EMP). The coefficients determine the impact of a one-standard deviation change in the predictors on the difference between the equity and option VRPs. The figures in parentheses report the t-statistics of the estimated coefficients computed using a bootstrap procedure described in the appendix. Panel B examines the incremental predictive power of the (standardized) broker-dealer variables. ** designates statistical significance at the 5% level.

	Default (DEF)	Ted Spread (TED)	R. Var. (RV)	PE Ratio (PE)	Inflation (PPI)	Employ. (EMP)
Default	0.24	-				
	(1.15)	-				
Default-All Variables	0.18	-	0.03	−0.14	0.15	0.26
	(0.77)	-	(0.13)	(0.83)	(1.02)	(1.52)
Ted Spread	-	−0.03				
	-	(−0.16)				
Ted Spread-All Variables	-	−0.07	0.18	−0.28**	0.18	0.18
	-	(−0.51)	(1.06)	(−2.04)	(1.30)	(1.32)
Together	0.18	−0.03				
	(0.97)	(−0.18)				
Together-All Variables	0.15	−0.10	0.11	−0.23	0.20	0.26
	(0.68)	(0.73)	(0.65)	(1.51)	(1.28)	(1.62)

Table 10: Implied Variance Regression

Panel A reports the estimated coefficients and the R^2 of a time-series regression of the quarterly implied variance (measured as the squared VIX index) on the set of macro-finance predictors that include the lagged realized variance (RV), the price/earnings ratio (PE), the default spread (DEF), the quarterly inflation rate (PPI), and the quarterly employment rate (EMP). The coefficients determine the impact of a one-standard deviation change in the predictors on the implied variance, and are computed using the GMM approach described in Section 2. The figures in parentheses report the t-statistics of the estimated coefficients that are robust to the presence of heteroskedasticity. Panel B examines the incremental predictive power of the orthogonalized broker-dealer variables. Panel B examines the incremental predictive power of the broker-dealer variables. The leftmost columns show the results for the broker-dealer leverage ratio (LEV) and the quarterly return of the prime broker index (PBI) in the presence of the macro-finance variables. The rightmost columns repeat the analysis after replacing the leverage ratio with the annual change in the leverage ratio (ΔLEV). ***, **, and * designate statistical significance at the 1%, 5%, and 10% levels, respectively.

Panel A: Macro-Finance Variables

	Mean	R. Var. (RV)	PE Ratio (PE)	Default (DEF)	Inflation (PPI)	Employ. (EMP)	R^2
R. Variance	1.23***	0.80***					0.72
	(23.19)	(9.62)					
All Variables	1.23***	0.76***	−0.12**	0.22***	−0.05	0.10	0.76
	(26.23)	(9.46)	(−2.01)	(3.40)	(−1.11)	(1.55)	

Panel B: Contribution of Broker-Dealer Variables

	Leverage (LEV)	PB Index (PBI)	R^2	ΔLeverage (ΔLEV)	PB Index (PBI)	R^2
+Leverage	−0.07*		0.77	−0.08**		0.77
	(−1.82)			(−2.14)		
+Prime Broker		−0.14**	0.78		-	-
		(−2.02)			-	
+Leverage & Prime Broker	−0.14***	−0.17**	0.79	−0.15***	−0.19***	0.79
	(−3.47)	(−2.42)		(−5.06)	(−2.85)	

Figure 1: Equity Variance Risk Premium

This figure reports the path of the quarterly equity Variance Risk Premium (VRP) obtained with the set of macro-finance predictors that includes the lagged realized variance, the price/earnings ratio, the default spread, the quarterly inflation rate, and the quarterly employment rate. The y-axis is in percent per quarter. Shaded areas correspond to NBER recession periods. Markers indicate the VRP for the quarter that follows the 1973 oil price shock (Oil Shock), the 1987 stock market crash (87 Crash), the beginning of the 1991 US military operation in Kuwait and Iraq (Gulf War), the 1994 bond sell-off after the sudden monetary tightening earlier the same year (Bond Sell-off), the 1998 collapse of the Long Term Capital Management fund (LTCM), the September 2001 terrorist attacks (9/11), the 2008 collapse of Lehman Brothers (Lehman), and the 2011 announcement of the Greek referendum on the exit from the Eurozone that followed the second rescue program (Greece).

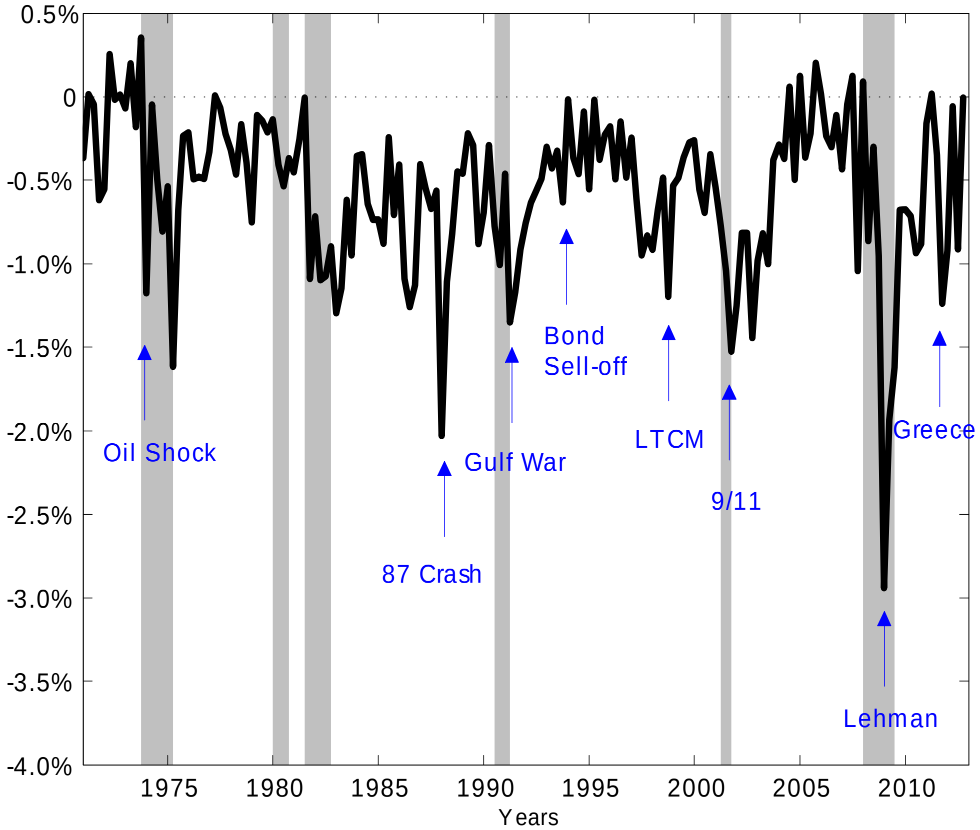

Figure 2: Option Variance Risk Premium

This figure reports the paths of the quarterly option Variance Risk Premium (VRP) obtained with the set of all predictors that includes the lagged realized variance, the price/earnings ratio, the default spread, the quarterly inflation rate, the quarterly employment rate, the broker-dealer leverage, and the quarterly return of the prime broker index. The path of the option VRP is only reported during the short sample (1992-2012) because the VIX index computed from three-month options is only available beginning in 1992. For comparison purposes, we also plot the equity VRP path previously depicted in Figure 2. The y-axis is in percent per quarter. Shaded areas correspond to NBER recession periods.

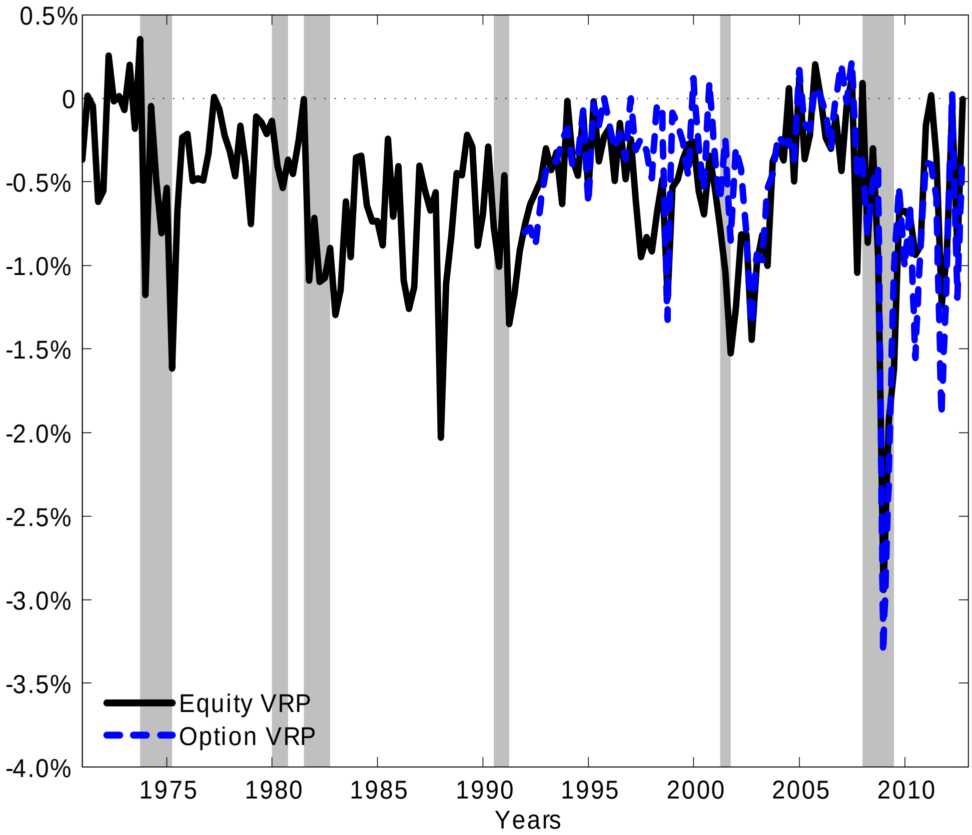

Figure 3: VRP Difference and Broker-Dealer Leverage

This figure plots the quarterly difference between the equity and the option VRP (black line). Each VRP is conditioned on the same set of predictors that includes the lagged realized variance, the price/earnings ratio, the default spread, the quarterly inflation rate, the quarterly employment rate, the broker-dealer leverage, and the quarterly return of the prime broker index. The dashed line shows the evolution of the quarterly leverage ratio of broker-dealers (in log form). The left y-axis is in percent per annum. Markers indicate the VRP difference for the quarter that the 1994 bond sell-off after the sudden monetary tightening earlier the same year (Bond Sell-off), the 1998 collapse of the Long Term Capital Management fund (LTCM), the September 2001 terrorist attacks (9/11), the 2008 collapse of Lehman Brothers (Lehman), and the 2011 announcement of the Greek referendum on the exit from the Eurozone that followed the second rescue program (Greece).

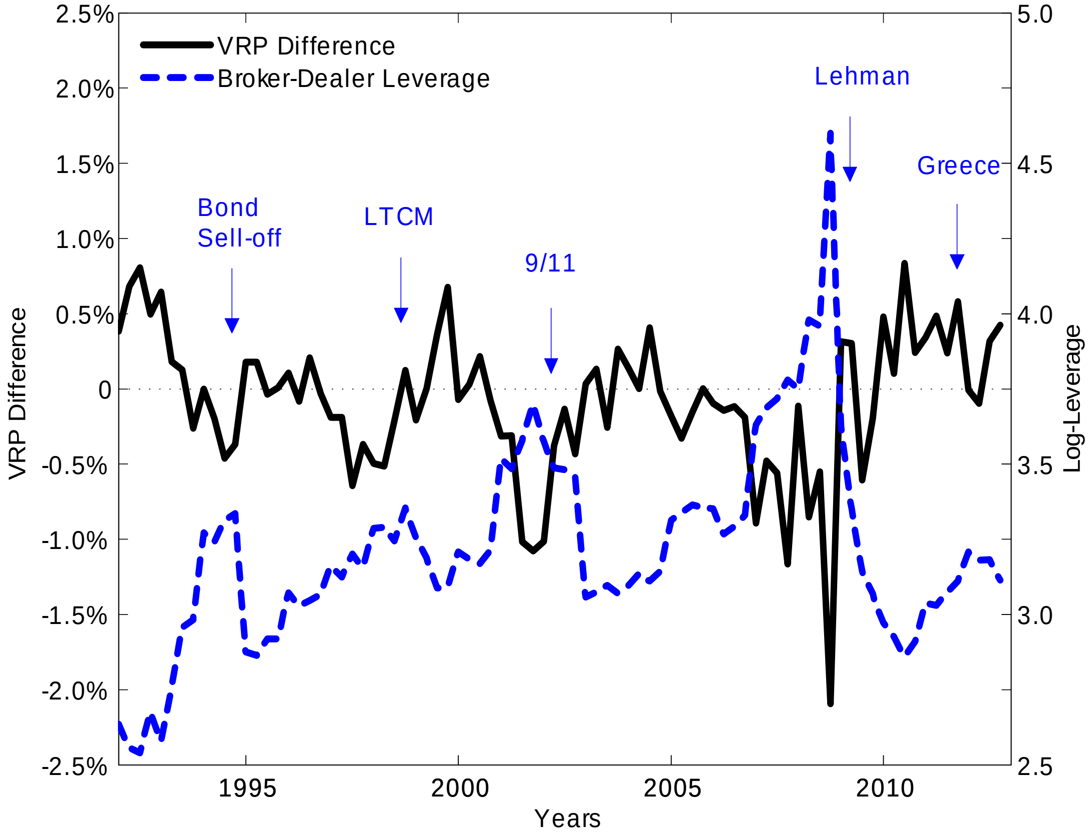

Figure 4: Mimicking Portfolio Innovations

This figure plots the innovations of the realized variance (red dotted line) and of the equity- (solid black line) and option-based (dashed blue line) factor mimicking portfolios. Innovations are conditioned on the same set of predictors that includes the lagged realized variance, the price/earnings ratio, the default spread, the quarterly inflation rate, the quarterly employment rate The y-axis is in percent per annum.