The Jump Leverage Risk Premium

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Abstract

Jumps in asset prices are ubiquitous, yet the apparent high price of jump risk observed empirically is commonly viewed as puzzling. We develop new model-free short-time risk-neutral variance expansions, allowing us to clearly delineate the importance of jumps in generating both price and variance risks. We find that simultaneous jumps in the price and the stochastic volatility and/or jump intensity of the market commands a sizeable risk premium. This empirically large “jump leverage” risk premium may be rationalized in the context of equilibrium-based models by jumps in the conditional moments of the underlying fundamentals and/or changes in investors’ risk aversion.

Keywords: jumps, options, tail risk, leverage effect, risk premiums, VIX index.

JEL classification: C51, C52, G12.

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1 Introduction

There is ample empirical evidence for the existence of jumps, or discontinuities, in the price of the market portfolio. There is also an extensive literature suggesting that the risk associated with market price jumps is priced differently from the risk associated with “smooth,” or continuous, price moves.\(^1\) Jumps also feature prominently in many equilibrium-based models seeking to rationalize the joint behavior of stock and option markets.\(^2\) These existing empirical analyses and theoretical models notwithstanding, the apparent “high” price of market jump risk is widely regarded as puzzling.

To more concretely demonstrate this apparent puzzle, consider the tail shape parameters of the actual and risk-neutral jump distributions that determine the rate at which the tails of the respective distributions decay to zero. Extreme Value Theory (EVT) dictates that this decay may be conveniently approximated by the tails of a generalized Pareto distribution. Relying on the S&P 500 high-frequency returns analyzed below, together with the nonparametric techniques of Bollerslev and Todorov (2011a), the resulting estimate for the observed left jump tail shape parameter equals 235. By comparison, the estimate for the risk-neutral left jump tail shape parameter based on the S&P 500 options data also analyzed below, together with the methods of Bollerslev and Todorov (2014) and Bollerslev et al. (2015), only equals 20. This large gap in the rate at which the two jump tails decay is difficult to rationalize with most existing economic models.\(^3\) As a case in point, for the time-varying disaster risk equilibrium models of Wachter (2013) and Seo and Wachter (2019), which are explicitly designed to account for the rare occurrence of market price jumps, this difference in the tail shape parameters would necessitate a coefficient of risk-aversion for the representative agent


\(^3\)This also echoes longstanding empirical evidence for the “expensiveness” of deep out-of-the-money put options; see, e.g., Bondarenko (2014b), among others.
in excess of two hundred.\footnote{This is in line with the recent study by Beason and Schreindorfer (2022), and the finding that most existing equilibrium-based models as traditionally calibrated are not consistent with where in the domain of the return space the equity premium is actually earned.}

Set against this background we develop new model-free short-time expansions for two alternative risk-neutral variance measures that allow us to more clearly delineate the nuanced effects of jumps and the pricing thereof.\footnote{Our paper is perhaps most closely related to the recent work by Orlowski et al. (2023), who measure the skewness risk premium through the average profits of a trading strategy that creates exposure to said risk in a model-free manner. Differentiating between trading and non-trading hours they find that much of the premium is earned overnight. Consistent with our findings they also find that during active trading hours the premium is dominated by priced jump risk, and that the premium tend to increase after left-tail events.} Our resulting risk premium estimates for the S&P 500 market index indicate a much more important role for “jump leverage effects,” or simultaneous jumps in the price and the stochastic volatility and/or jump intensity of the market, than hitherto suggested in the literature. Correspondingly, the compensation for jump leverage risk also accounts for a sizeable portion of the much-studied market variance risk premium.\footnote{Hu et al. (2022) have also recently documented a strong positive empirical relation between the variance risk premium and the conditional covariance between the market return and its variance, or the “leverage effect.” The discrete-time stochastic volatility model in Cheng et al. (2019) also relies on a related identification scheme.} In short, the puzzling “high” price of market jump risk may be explained by the fact that this risk is typically accompanied by other risks that are similarly disliked by investors.

To help intuit the key empirical features underlying these findings, the left panel in Figure 1 shows the intraday prices for the S&P 500 market index at a 5-minute frequency for the week of March 4 - 11, 2020. Due to heightened fears about the global pandemic, the index dropped by more than 10% for that week as a whole. Meanwhile, even though the single large overnight decline on March 6 clearly stands out, there were also many other smaller intraday price jumps during that week. Most of these smaller price jumps also appeared to be followed by heightened volatility. Corroborating that idea, the right panel in Figure 1 shows that while the realized jump risk for the week as a whole (indicated by the dashed line at week 0) did indeed increase compared to the week before (indicated by the dashed line at week -1), so did the total realized volatility inclusive of the diffusive price risk (indicated by the solid line).\footnote{The total weekly realized variation measures are based on the summation of the squared 5-minute returns} In other words, negative jumps in the market index tend
Figure 1: The left panel shows the SPY at a 5-minute intraday frequency for the week of March 4 - 11, 2020. The right plot shows the total weekly realized variation based on the summation of the squared 5-minute returns (solid line) and the realized jump variation (dashed line) constructed as the difference between the total variation and the bipower variation. The week of March 4 - 11, 2020 corresponds to week 0. Both of the variation measures are reported in annualized volatility units.

to induce positive jumps in the aggregate market volatility and the intensity of future jump arrivals.\(^8\)

We are not the first to explicitly highlight the existence of such a jump leverage effect, or the tendency for negative market price jumps to trigger positive jumps in market volatility and/or market jump intensity.\(^9\) However, the pricing implications of this effect and its role in explaining the seemingly “high” price of market jump risk has not previously been thoroughly studied in the literature. We seek to fill this void and precise the impact of the

\(^8\)These observations are also broadly consistent with the high-frequency-based parametric model estimates in Bates (2019), and the finding that large daily market moves typically represent the accumulation of a series of self-exciting intradaily volatility-price cojumps; see also the more recent estimates in Ewald and Zou (2021).

\(^9\)See e.g., Jacod and Todorov (2010), Todorov and Tauchen (2011), Bandi and Renò (2012), Bandi and Renò (2016), Ait-Sahalia et al. (2017), Jacod et al. (2017), among others. There is also a large existing literature concerned with the nonparametric estimation of leverage effects more generally without separately considering the impact of jumps; see e.g., Bollerslev et al. (2006), Ait-Sahalia et al. (2013), Wang and Mykland (2014), Andersen et al. (2015) and Kalnina and Xiu (2017), among others.
jump leverage effect for explaining market jump risk premiums. More specifically, utilizing the rich information in short-dated options our analysis clearly shows that volatility jumps and price-volatility co-jumps are not merely modeling choices, but are a necessity for properly understanding how jump risk actually propagates and how it is priced.\textsuperscript{10}

Our approach is decidedly non-parametric, relying on the rich information in short-dated options with different tenors for extracting information about the pricing of the jump leverage effect.\textsuperscript{11} Intuitively, while asymmetry in the jump distribution and the jump leverage effect will both contribute to asymmetry in the distributions of returns over very short horizons, it is possible to separate the two effects by also looking at slightly longer horizons and the term structure of skewness.\textsuperscript{12} In particular, while the asymmetry of jump risk should manifest the same across different horizons, the leverage effect should feature more prominently over longer return horizons. As such, our use of extremely short-dated options combined with our nonparametric techniques allows us to more precisely identify how the jumps propagate and how the corresponding short-term risks are priced. Importantly, we do not seek to explicitly model the dynamics of the corresponding spot volatility and jump intensities of the price jumps and the price-volatility co-jumps, as would be required to study the propagation of risks over longer horizons, instead we rely solely on the non-parametric model-free procedures and focus on the short-term pricing. This also clearly distinguishes our study from other existing work.

To illustrate, Figure 2 plots the (normalized to unity at the shortest horizon) term structures of the risk-neutral variance and risk-neutral third moment calculated from short-dated S&P 500 options on two select days: February 4, 2019, a day with relatively low volatility, and March 26, 2020, a day with relatively high volatility.\textsuperscript{13} Since volatility is well-known to be mean reverting, the risk-neutral variance term structure is naturally upward sloping on

\textsuperscript{10}By contrast, earlier work based on the estimation of parametric stochastic volatility models and options with longer tenors than the ones used here often report much smaller economic gains from incorporating volatility jumps, and jump leverage effects, in the price dynamics; see e.g., Broadie et al. (2007).

\textsuperscript{11}Jackwerth and Vilkov (2019) have also recently proposed an estimator for the risk-neutral leverage effect based on a parametric copula model and options on the S&P 500 and VIX index. By contrast, our method is fully nonparametric and seeks to estimate the instantaneous jump leverage effect.

\textsuperscript{12}Note, this is distinctly different from the realized skewness of returns over long horizons analyzed in the work by Neuberger (2012).

\textsuperscript{13}The estimates for the variance and the third moment shown in the figure are based on the expressions for $\hat{\nu}^Q_t,T$ and $3(\hat{\nu}^Q_t,T - \hat{\nu}^0_t,T)$ formally defined in equations (21) and (22) below.
Figure 2: The left two panels show the second risk neutral moments calculated from short-dated S&P 500 options for different horizons on February 4, 2019 (top) and March 26, 2020 (bottom). The right two panels show the third risk neutral moments on the same two days. For easy of comparisons, the moments at the shortest horizons are normalized to unity.

The low volatility day and downward sloping on the high volatility day. By contrast, the term structure for the risk-neutral third moment is strongly upward sloping on both days. Looking across other high and low volatility days reveal the same general pattern: the risk-neutral third moment term structure is almost always upward sloping. This finding is difficult to reconcile with mean reversion in state variables, including the volatility and/or jump intensity, as one would expect these effects to manifest in opposite directions on high and low volatility days. Instead, the systematic upward sloping term structure for the risk-neutral third moment points to the significant pricing of the jump leverage effect, with the steepness of the slopes further underscoring the seemingly large magnitude of the premium.\textsuperscript{14}

\textsuperscript{14}Relatedly, Bakshi et al. (2003), Kozhan et al. (2013) and Schneider et al. (2020) have previously emphasized the nontrivial pricing of skewness and coskewness risk embedded in options. The risk-neutral third moment is, of course, also formally affected by the total leverage effect; i.e., the covariation between price and volatility shocks in general. Meanwhile, as argued below it appears impossible to reconcile the empirically very steep short-term slope observed in Figure 2 without explicitly allowing for significant pricing of the jump leverage effect. The additional results for the parametric double-jump stochastic volatility model
In order to more formally assess this pricing, we begin our analysis by nonparametrically identifying the risk-neutral expectation of the jump leverage effect. We do so by developing new short-term expansions for two alternative risk-neutral variance measures: (i) the risk-neutral conditional expectation of the log-price, mirroring the measure used by the Chicago Board Options Exchange (CBOE) in their computation of the VIX index, and (ii) the risk-neutral conditional log-return variance. While both of our expansions are linear functions of the return time horizon and its square, with the leading term in both being equal to the risk-neutral spot variance, the slopes of the two expansions differ. Moreover, as formally demonstrated below, this difference is almost exclusively attributable to the jump leverage effect, allowing for its nonparametric estimation from actively traded short-dated S&P 500 index options. Comparing the resulting risk-neutral estimates with their realized counterparts estimated from high-frequency intraday S&P 500 index returns, in turn reveal highly significant compensation for short-term leverage risk, with more than half of the risk premium in third return moments at a horizon of two weeks attributable to jump leverage.

From a theoretical perspective, this significant pricing of the jump leverage effect is naturally associated with instantaneous changes in the investment opportunity set induced by jumps in the volatility and/or jump intensity. Accordingly, shocks to the stochastic volatility and/or jump intensity should be priced. Consistent with this thesis, we find that at a horizon of two weeks, close to one-fifth of the variance risk premium may be directly attributed to compensation for time variation in the stochastic volatility and/or jump intensity, as opposed to compensation for instantaneous variance risk. Further corroborating the non-trivial pricing of the jump leverage effect, we also find that the term structure of the variance risk premium substantially flattens for horizons beyond two weeks.\textsuperscript{15}

In addition to our main empirical findings based on the pricing of short-dated S&P 500 index options, we also provide additional supportive empirical evidence based on the pricing of VIX index options. In particular, while the realized jump risk for the VIX index appears close to symmetric, we find that its risk-neutral counterpart implied from VIX index options and our new short-term expansions for the option-implied variance risk measures is highly
discussed in Appendix C also further corroborate this.
\textsuperscript{15}This flattening of the term structure over longer horizons is also in line with the empirical evidence of Dew-Becker et al. (2017).
right skewed, as would be the case if jump leverage risk is significantly priced.

The rest of the paper is organized as follows. We begin in Section 2 by formally developing the short-time expansions for the risk-neutral variance measures that underly our analyses. Sections 3 and 4, respectively, define the different jump and leverage risk premium measures and discusses their practical implementation based on short-dated S&P 500 options and high-frequency intraday S&P 500 returns. Section 5 presents our main new empirical findings pertaining to the pricing of jump leverage risks. Section 6 provides additional supportive empirical evidence based on the pricing of VIX index options. We conclude with a brief informal discussion of various equilibrium-based models that can possibly help explain the nontrivial pricing of the jump leverage effect. All of the proofs, along with Monte Carlo simulation evidence underscoring the accuracy of the new variance expansions, are deferred to an Appendix.

2 Short-Time Expansions of Risk-Neutral Variance Measures

We will consider two alternative risk-neutral variance measures and corresponding expansions. The difference between the two expansions in turn allow us to identify and nonparametrically estimate the risk-neutral jump leverage effect from short-dated options.

Let $X_t$ denote the underlying price process for the market, with the corresponding log-price process denoted $x_t$. We will assume that $X_t$ is defined on the probability space $(\Omega, \mathcal{F}, (\mathcal{F}_t)_{t\geq 0}, \mathbb{P})$, and that the price obeys the following general Itô semimartingale dynamics

$$\frac{dX_t}{X_{t-}} = \alpha_t dt + \sigma_t dW_t + \int_{\mathbb{R}} (e^{z} - 1)(\mu - \nu^P)(dt,dz),$$

where $W_t$ is a Brownian motion, $\mu$ is an integer-valued jump measure counting the jumps in $X$ with compensator $\nu^P$, $\alpha_t$ denotes the drift capturing the instantaneous expected return, and $\sigma_t$ refers to the instantaneous stochastic volatility. Under standard conditions, no-arbitrage implies the existence of a risk-neutral measure under which the cum-dividend discounted price process is a local martingale. Since we will be focusing on short return horizons, for
notational simplicity we will set the risk-free rate and the dividend yield identically equal to zero.\footnote{The extension to allow for nonzero interest rate and/or dividend yield is discussed in Appendix B. This appendix also shows that the effect from ignoring the risk-free interest rate and the dividend yield in our analysis is small.} Accordingly, the dynamics of $X_t$ under the risk-neutral probability measure $\mathbb{Q}$ may be expressed as

$$
\frac{dX_t}{X_{t-}} = \sigma_t dW_t^\mathbb{Q} + \int_{\mathbb{R}} (e^z - 1)(\mu - \nu^\mathbb{Q}) (dt, dz),
$$

(2)

where $W_t^\mathbb{Q}$ is a $\mathbb{Q}$-Brownian motion and $\nu^\mathbb{Q}$ denotes the jump compensator under $\mathbb{Q}$.

Our first risk-neutral variance measure is simply given by the conditional variance of the log-return over some time horizon $T$ under the above defined $\mathbb{Q}$ measure. That is:

$$
V_{t,t+T}^\mathbb{Q} \equiv \frac{1}{T} \text{Var}_t^\mathbb{Q} (x_{t+T} - x_t).
$$

(3)

Our second variance measure is based on the conditional risk-neutral mean of the log-return over the horizon $T$,

$$
\mathcal{V}_{t,t+T}^\mathbb{Q} \equiv -\frac{2}{T} \mathbb{E}_t^\mathbb{Q} (x_{t+T} - x_t) = \frac{1}{T} \mathbb{E}_t^\mathbb{Q} \left( \int_t^{t+T} \sigma_s^2 ds + 2 \int_t^{t+T} \int_{\mathbb{R}} (e^z - 1 - z) \nu^\mathbb{Q}_s (ds, dz) \right),
$$

(4)

where the second equality follows by an application of Itô’s lemma. This second variance measure is also the measure used by the CBOE in the construction of the VIX volatility index.\footnote{The connection between the $V_{t,t+T}^\mathbb{Q}$ and $\mathcal{V}_{t,t+T}^\mathbb{Q}$ variance measures has also previously been used by Du and Kapadia (2012) in the construction of an options-based jump index, while Bondarenko (2014a) has previously analyzed the connection between the two measures in regards to variance trading and replication strategies.}

Letting the return horizon approach zero, the instantaneous counterparts of the above two variance measures are naturally defined by:

$$
V_t^\mathbb{Q} \equiv \lim_{\Delta \to 0} V_{t,t+\Delta}^\mathbb{Q}, \quad \mathcal{V}_t^\mathbb{Q} \equiv \lim_{\Delta \to 0} \mathcal{V}_{t,t+\Delta}^\mathbb{Q}.
$$

(5)

These instantaneous measures may also alternatively be expressed in terms of the risk-neutral
asset characteristics as

$$V_t^Q = \sigma_t^2 + \int_R z^2 \nu_t^Q(dz), \quad V_t^Q = \sigma_t^2 + 2 \int_R (e^z - 1 - z) \nu_t^Q(dz),$$  \hspace{1cm} (6)$$

where the risk-neutral jump compensator is decomposed as $\nu^Q(dt, dz) = dt \nu_t^Q(dz)$. As these expressions make clear, both of the instantaneous variance measures only depend on the $Q$ probability measure through their second jump components. As such, the difference between the two measures may be entirely explained by jumps,

$$V_t^Q - V_t^Q = \int_R z^2 \nu_t^Q(dz) - 2 \int_R (e^z - 1 - z) \nu_t^Q(dz) \approx -\frac{1}{3} \int_R z^3 \nu_t^Q(dz),$$  \hspace{1cm} (7)$$

that is the (scaled) $Q$ third moment of market price jumps.

Over ultra short return horizons, or $T \approx 0$, the $V_{t,t+T}^Q$ and $V_{t,t+T}^Q$ variance measures may naturally be used as proxies for the instantaneous variance measures $V_t^Q$ and $V_t^Q$, respectively. However, for longer return horizons, or $T > 0$, mean reversion in volatility and jump intensity will both contribute nontrivially to $V_{t,t+T}^Q$ and $V_{t,t+T}^Q$, thus rendering the difference between the two measures more difficult to interpret. In order to formally account for this, it is useful to define the instantaneous drift term of any arbitrary process $z$ evaluated under $Q$ as:

$$m_t^Q(z) \equiv \lim_{\Delta \to 0} \frac{\mathbb{E}_t^Q(z_{t+\Delta} - z_t)}{\Delta}. \hspace{1cm} (8)$$

Of course, if the process $z$ is stationary, then $\mathbb{E}_t^Q(m_t^Q(z)) = 0$. Meanwhile, any risk premium for variation in $z$ will generally imply that $\mathbb{E}_t^P(m_t^Q(z)) \neq 0$, with the sign being positive (negative) if the specific variation in $z$ is disliked (liked) by investors.

Now, utilizing the above notation, it is possible to show that

$$V_{t,t+T}^Q = V_t^Q + \frac{T}{2}\times m_t^Q(V_t^Q) + T^2 \times C_t + O_p(T^3), \quad T \downarrow 0, \hspace{1cm} (9)$$

where $C_t$ denotes some $\mathcal{F}_t$-adapted random variable.\footnote{Formal proofs of this expansion and the expansion presented below are provided in Appendix A.} Notably, the leading term in the expansion, which does not depend upon $T$, is given by the instantaneous risk-neutral variance
while the slope of the expansion, when viewed as a function of $T$, is solely determined by the mean reversion of $V_t^Q$. The random variable $C_t$ that dictates the $T^2$ term depends in a complicated way on various features of the volatility and jump intensity dynamics. However, this term is numerically small, and we will treat it as a nuisance parameter in what follows.

To establish an analogous expansion for the $V_{t,t+T}^Q$ variance measure, it is instructive to first formally define the notion of an instantaneous leverage effect, that is a measure of the co-dependence between innovations, or shocks, to the market price and return variation. To this end, for arbitrary processes $y$ and $z$, denote their quadratic covariation by:

$$[y, z]_t = \lim_{n \to \infty} \sum_{i=1}^{nt} [(y_{i/n} - y_{(i-1)/n})(z_{i/n} - z_{(i-1)/n})].$$

Furthermore, denote the predictable counterpart of $[y, z]_t$ under the $Q$ measure by $\langle y, z \rangle_t^Q$,

$$[y, z]_t = \langle y, z \rangle_t^Q + Q\text{-martingale}.$$

In parallel to the instantaneous variance and drift measures defined in (5) and (8), respectively, the instantaneous risk-neutral leverage effect is then naturally defined by:

$$L_t^Q \equiv \lim_{\Delta \to 0} \frac{\langle x, V_t^Q \rangle_{t+\Delta} - \langle x, V_t^Q \rangle_t}{\Delta}.$$  

Using this definition of $L_t^Q$, the following short-term expansion obtains

$$V_{t,t+T}^Q = V_t^Q + \frac{T}{2} \times (m_t^Q(V_t^Q) - L_t^Q) + T^2 \times C_t + O_p(T^3), \quad T \downarrow 0,$$

where again $C_t$ denotes some numerically small $\mathcal{F}_t$-adapted random variable.

The expressions for $V_{t,t+T}^Q$ and $V_{t,t+T}^Q$ in (9) and (13) are obviously very similar except for one important distinction, namely the presence of the $L_t^Q$ leverage term in the slope of the latter expansion. There are at least two intuitive reasons for this important “extra” term, representing the co-movements between the conditional mean and the martingale component of the log-price. Firstly, $V_{t,t+T}^Q$ is defined as a centered second moment, while $V_{t,t+T}^Q$ is not. Secondly, $L_t^Q$ is computed under the $Q$ measure, and as discussed more formally in Appendix
A, considerations of no-arbitrage explicitly restricts the drift term under the risk-neutral measure to be a function of the jump intensity and the diffusive volatility. By comparison, there is no such restriction under the $\mathbb{P}$ measure, which means that a short-time expansion of the $\mathbb{P}$ counterpart to $V_{t,t+T}^Q$ would also look different.

Meanwhile, since the leverage effect and the dependence between price and variance innovations is generally found to be negative, the $L_t^Q$ term will tend to increase the term structure slope of $V_{t,t+T}^Q$ compared to the slope of $\mathcal{V}_{t,t+T}$. In fact, since $m_t^Q(V_t^Q) - m_t^Q(V_t^Q)$ is likely to be small in absolute value, as it depends on the third moment of jumps, the term structure slope of $V_{t,t+T}^Q - V_{t,t+T}^Q$ will predominantly be determined by the $-L_t^Q$ leverage term. As such, this helps explain the seemingly puzzling empirical evidence discussed in the introduction, and the systematically upward sloping term structure of the third risk neutral moment irrespective of the level of the volatility. We will return to this in our discussion of the estimated jump leverage risk premium below.

3 Jump and Leverage Risk Premiums

The risk-neutral measures defined in connection with the expansions discussed in the previous section naturally suggest the definition of corresponding risk premiums to help illuminate the pricing of jump risk, and jump leverage risk in particular.\footnote{To help fix ideas, Appendix C further discusses how the different risk premiums defined below manifest in the context of the parametric double-jump stochastic volatility model of Duffie et al. (2000).}

To begin, the instantaneous variance risk premiums for the two different variance measures may simply be defined by

$$ IVRP_t \equiv V_t^Q - V_t^P, \quad TVRP_t \equiv \mathcal{V}_t^Q - \mathcal{V}_t^P, $$

(14)

where $V_t^P$ and $\mathcal{V}_t^P$ denote the $\mathbb{P}$ counterparts to the $V_t^Q$ and $\mathcal{V}_t^Q$ instantaneous variances. Since the diffusive spot volatility $\sigma_t^2$ must be identical under the $\mathbb{P}$ and $\mathbb{Q}$ measures in order to prevent arbitrage opportunities, the $IVRP_t$ and $TVRP_t$ risk premiums are solely determined by the pricing of jumps. However, that is not the case for the corresponding variance risk
premiums defined over non-trivial time intervals $T > 0$,

$$VRP_{t,t+T} \equiv V^Q_{t,t+T} - V^P_{t,t+T}, \quad VRP_{t,t+T} \equiv V^Q_{t,t+T} - V^P_{t,t+T}. \quad (15)$$

The second of these two measures, in particular, has been extensively studied empirically in the recent literature, typically with $T$ set to one-month mimicking the horizon of the popular VIX index used in place of $V^Q_{t,t+T}$.\(^{20}\)

To help more formally assess what drives the $VRP_{t,t+T}$ and $VRP_{t,t+T}$ measures over short horizons $T$, it is instructive to employ the expansions developed in the previous section. Doing so, we obtain the following two decompositions:

$$VRP_{t,t+T} = IVRP_t + \frac{T}{2} \times (m^Q_t(V^Q_t) - m^P_t(V^Q_t)) - \frac{T}{2} \times (L^Q_t - L^P_t) + O_p(T^2), \quad (16)$$

and

$$VRP_{t,t+T} = IVRP_t + \frac{T}{2} \times (m^Q_t(V^Q_t) - m^P_t(V^Q_t)) + O_p(T^2). \quad (17)$$

As previously noted, the leading component in $VRP_{t,t+T}$ (resp. $VRP_{t,t+T}$), that is the instantaneous variance risk premium $IVRP_t$ (resp. $IVRP_t$), is solely determined by asset price jumps. On the other hand, the second $m^Q_t(V^Q_t) - m^P_t(V^Q_t)$ (resp. $m^Q_t(V^Q_t) - m^P_t(V^Q_t)$) term, which depends linearly on the horizon $T$, reflects compensation demanded by investors for any changes in the investment opportunity set. In the context of the underlying general Itô semimartingale in (1), and its risk-neutral counterpart in (2), that is temporal variation in $\sigma_t^2$ and/or changes in the intensity of the jumps.

Importantly, the decomposition for $VRP_{t,t+T}$ contains an additional linear-in-$T$ component specifically due to the compensation for leverage risk. We will refer to this “extra” term as the instantaneous leverage risk premium in the sequel, or

$$ILRP_t \equiv L^Q_t - L^P_t. \quad (18)$$

In parallel to the instantaneous variance risk premiums, which are solely determined by

\(^{20}\)See, e.g., Bollerslev et al. (2009), Drechsler and Yaron (2011) and Bekaert and Hoerova (2014), along with many subsequent studies.
jumps and the pricing thereof, $ILRP_t$ is solely determined by the jump leverage effect and the pricing thereof, that is the risk premium for co-jumps between the log-price $x$ and $V^Q$. Empirically, we would expect $ILRP_t$ to be negative and thus contribute positively to $VRP_{t,t+T}$.

Note that even though $ILRP_t$ does not appear directly in $VRP_{t,t+T}$, that does not mean that $VRP_{t,t+T}$ is not formally affected by the jump leverage effect. Indeed, if price-volatility co-jumps are priced, then this price will also formally manifest in $m_t^Q(V_t^Q) - m_t^P(V_t^Q)$ (as well as in $m_t^Q(V_t^Q) - m_t^P(V_t^Q)$). This is also further illustrated in the context of the popular parametric double-jump stochastic volatility model of Duffie et al. (2000) discussed in Appendix C. Relatedly, it is somewhat misleading to judge the overall importance of the jump leverage effect for the variance risk premium by simply comparing the magnitude of $ILRP_t$ with that of $VRP_{t,t+T}$. Of course, $ILRP_t$ and $VRP_{t,t+T}$ are also measured in “units” of third and second moments, respectively, analogous to $V_t^Q - V_t^Q$ defined in (7) and the spot volatility $V_t^Q$.

We turn next to a discussion of our model-free estimation of these different risk premiums.

4 Feasible Measures

Our feasible counterparts to the risk premium quantities defined in the previous section rely on portfolios of short-dated options to estimate the $Q$ risk-neutral quantities and high-frequency returns to proxy the corresponding $P$ risk measures.

4.1 Option-Based Risk Measures

Let $O_{t,T}(K)$ denote the time $t$ price of a European-style out-of-the-money option expiring at time $t+T$ with strike $K$. With the risk-free rate and the dividend yield both set to zero, the general results in Bakshi and Madan (2000) and Carr and Madan (2001) then provide the following two option spanning results:

$$
E_t^Q(x_{t+T} - x_t)^2 = 2 \int_0^\infty \left( 1 - \log \left( \frac{K}{X_t} \right) \right) \frac{O_{t,T}(K)}{K^2} dK,
$$

(19)
and
\[ \mathbb{E}_t^Q (x_{t+T} - x_t) = - \int_0^\infty \frac{O_{t,T}(K)}{K^2} dK. \] 
(20)

In practice, of course, we do not observe options on a continuum of strikes. Instead let the discrete grid of \( N_{t,T} \) option prices observed at time \( t \) be denoted by \( K_1 < \ldots < K_{N_{t,T}} \), where for simplicity we suppress the dependence of the strike grid on the pair \((t, T)\). The actually observed option prices \( \hat{O}_{t,T}(K_j) \) are also subject to pricing errors, say \( \hat{O}_{t,T}(K_j) = O_{t,T}(K_j) + \epsilon_{t,T}(j) \) for \( j = 1, \ldots, N_{t,T} \). Following standard practice in the option pricing literature, we will assume that the \( \epsilon_{t,T}(j) \) observation errors exhibit only weak spatial and temporal dependencies, and hence are “averaged out” in the estimation.

Using the observed option prices and the spanning results in (19) and (20), we construct the following estimates for \( V_{t,t+T}^Q \) and \( \hat{V}_{t,t+T}^Q \):

\[ \hat{V}_{t,t+T}^Q = \frac{2}{T} \sum_{j=2}^{N_{t,T}} \left( 1 - \log \left( \frac{K_{j-1}}{X_t} \right) \right) \frac{\hat{O}_{t,T}(K_{j-1})}{K_{j-1}^2} (K_j - K_{j-1}) - \frac{T}{4} \left( \hat{V}_{t,t+T}^Q \right)^2, \] 
(21)

\[ \hat{V}_{t,t+T}^Q = \frac{2}{T} \sum_{j=2}^{N_{t,T}} \hat{O}_{t,T}(K_{j-1}) \frac{K_{j-1}^2}{K_j^2} (K_j - K_{j-1}). \] 
(22)

Guided by the expansions for the risk-neutral variance measures in (9) and (13), we then run the following linear regressions at each point time using all of the available tenors:

\[ \hat{V}_{t,t+T}^Q = b_{t,0} + b_{t,1} T_j + b_{t,2} T_j^2 + \epsilon_{t,T_j}, \quad \hat{V}_{t,t+T}^Q = \beta_{t,0} + \beta_{t,1} T_j + \beta_{t,2} T_j^2 + \epsilon_{t,T_j}. \] 
(23)

Denoting the resulting OLS estimates by \( \hat{b}_{t,i} \) and \( \hat{\beta}_{t,i} \), respectively, our risk-neutral variance estimates are then simply defined by:

\[ \hat{V}_t^Q = \hat{b}_{t,0}, \quad \hat{V}_t^Q = \hat{\beta}_{t,0}. \] 
(24)

In lieu of (9) and (13), \( L_t^Q \) may seemingly be estimated by twice the difference in the estimated slopes from the two regressions in (23). However, this estimator will be biased due to the presence of the risk-neutral instantaneous drift \( m_t^Q (V_t^Q - V_t^Q) \) and the mean reversion
in \( \int_{\mathbb{R}} (z^2 - 2(e^z - 1 - z)) \nu_t^Q (dz) \). Even though this bias will typically be much smaller than \( \int_{\mathbb{R}} z^2 \nu_t^Q (dz) \), if the mean reversion in the jump intensity and the diffusive volatility manifest in the same direction, we can easily correct for the bias. In particular, utilizing

\[
\hat{s}_t = \frac{\sum_{i=1}^{k_n} |\hat{V}^Q_{t-i/n} - \hat{V}^Q_{t-(i+1)/n}|}{\sum_{i=1}^{k_n} |\hat{V}^{\overline{Q}}_{t-i/n} - \hat{V}^{\overline{Q}}_{t-(i+1)/n}|},
\]

based on a “small” window of \( k_n \) variance increments, we obtain the following simple bias-corrected leverage estimator:

\[
\hat{L}_t^Q = -2 \left( \hat{b}_{t,1} - \hat{s}_t \hat{\beta}_{t,1} \right).
\]

The Monte Carlo simulation results reported in Appendix D underscore the accuracy of each of the \( \hat{V}_t^Q \), \( \hat{V}_t^{\overline{Q}} \) and \( \hat{L}_t^Q \) risk-neutral risk estimators.

We turn next to a discussion of our corresponding return-based risk estimates.

### 4.2 Return-Based Risk Measures

We assume that market prices and options data are sampled \( n \) times during the unit time interval. Relying on a window consisting of \( k_n \) return and variance increments prior to time \( t \), we define our realized variance and leverage estimators as

\[
\hat{RV}_t = \frac{n}{k_n} \sum_{i=1}^{k_n} (x_{t-i/n} - x_{t-(i+1)/n})^2,
\]

and

\[
\hat{RL}_t = \frac{n}{k_n} \sum_{i=1}^{k_n} [(x_{t-i/n} - x_{t-(i+1)/n})(\hat{V}^Q_{t-i/n} - \hat{V}^Q_{t-(i+1)/n})],
\]

respectively. Under standard conditions, \( \hat{RV}_t \) and \( \hat{RL}_t \) consistently, for \( n \to \infty \), estimate \( QV_{t-k_n/n,t} \) and \( QL_{t-k_n/n,t} \), formally defined as the quadratic variation of \( x \) and the quadratic covariation of \( x \) and \( V^Q \),

\[
QV_{t-k_n/n,t} \equiv \frac{n}{k_n} ([x, x]_t - [x, x]_{t-k_n/n}),
\]

15
and
\[ QL_{t-k_n/n,t} = \frac{n}{k_n} ([x, \mathcal{V}_t^Q]_t - [x, \mathcal{V}_t^Q]_{t-k_n/n}). \] (30)

Moreover, if \(k_n \to \infty\) and \(k_n/n \to 0\), the \(QV_{t-k_n/n,t}\) and \(QL_{t-k_n/n,t}\) estimands naturally converge to their spot counterparts.

4.3 Feasible Realized Risk Premium Estimates

Armed with the above risk-neutral and realized risk estimates, our feasible versions of \(IVRP_t\) and \(ILRP_t\), defined in (14) and (18), are simply obtained as
\[ \hat{RIVRP}_t \equiv \hat{V}_t^Q - \hat{RV}_t, \] (31)
and
\[ \hat{RILRP}_t \equiv \hat{L}_t^Q - \hat{RL}_t, \] (32)
respectively. These estimators of the instantaneous variance and leverage risk premiums purposely rely on the easy-to-calculate model-free realized \(\hat{RV}_t\) and \(\hat{RL}_t\) measures, rather than their conditional expectations, \(\mathcal{V}_t^P\) and \(L_t^P\). Practical estimation of the conditional expectations would necessitate additional modeling assumptions pertaining to the dynamics of \(\mathcal{V}_t^P\) and \(L_t^P\). Importantly, however, relying on the realized values instead of their conditional expectations, does not alter any of our main conclusions and average risk premium estimates. In particular, ignoring the (negligible) contribution stemming from estimation error, it readily follows that
\[ \hat{RIVRP}_t = IVRP_t + \mathbb{P}\text{-martingale}, \quad \hat{RILRP}_t = ILRP_t + \mathbb{P}\text{-martingale}, \]
where the \(\mathbb{P}\)-martingale terms reflect the differences between the respective realizations of the discrete-time stochastic processes and their conditional expectations. Although these \(\mathbb{P}\)-martingale terms are not necessarily “small” numerically, they are by definition mean-zero and unpredictable, and hence do not affect the expected values, nor the dynamic dependencies, of the estimated risk premiums.
5 Market Jump and Leverage Risk Premiums

We begin the discussion of our main empirical findings with an account of the data underlying our results, followed by a summary of the daily risk estimates, before finally considering the implications of our model-free estimates for the pricing of jumps and jump leverage risk.

5.1 Data

Our empirical analyses is based on high-frequency returns and option data for the S&P 500 index spanning the period 2007-2020. The options are European style and are traded on the CBOE. We obtained the option data from the CBOE Data Shop. We rely on high-frequency price records of the SPY exchange traded fund (ETF) designed to track the S&P 500 index. The SPY data is obtained from the TAQ database. We record the SPY and options prices at a 5-minute frequency during the trading day, starting at 9.35 EST and ending at 15.55 EST, resulting in a total of 77 price records per day.

We apply standard filters and cleaning procedures to the data. In so doing, we remove any days for which the number of SPY zero returns exceeds twenty percent of the total daily number of high-frequency returns. This mostly eliminates trading days around holidays. We also remove the four 15-minute periods on March 9, 12, 16 and 18, 2020, where trading was halted due to market-wide circuit breakers. We take the option mid-quotes as our option price observations. We remove any options with zero bids and options for which the ratio of the ask relative to the bid exceeds ten. We also remove date and maturity pairs for which the minimum of the ratio of the out-of-the-money option price relative to its strike exceeds five percent of the maximum of this ratio. We further remove date and maturity pairs for which the maximum strike gap around the money exceeds twenty. Finally, to avoid anomalous results associated with large event risks, we also exclude any options for which the horizons span the dates of the 2016 Brexit and 2020 U.S. Presidential elections.

For the actual estimation, at each point in time we use the first two available shortest maturity options with at least three business days to expiration. We then keep adding options with tenors up until twelve business days. If on a given day the number of available tenors is just two, or if the gap between the shortest and longest available tenor is less than
six business days, we drop the squared tenor term in the regressions in (23) in order to avoid any issues with multicollinearity (this mostly happens in the early part of the sample). Also, if at a given point in time, any of the estimated intercept terms in (23) are negative, we simply remove those observations from the analysis (again this rarely happens, and mostly so around holidays).

The above choice of option tenors aims at striking a balance between bias and variance in the estimation. Since the expansion results in (9) and (13) explicitly pertain to small \( T \), to reduce the bias ideally we would like to only use the shortest possible tenors on a given day. However, restricting the estimation to only the very shortest maturities will obviously limit the number of options, in turn resulting in noisier estimates. Motivated by the Monte Carlo simulation results discussed in Appendix D, and the finding that even for very high levels of mean reversion, the biases in the estimation for tenors up to around twelve business days appear quite small, we deliberately include options with maturities up to that horizon.

5.2 Risk Estimates

We begin by computing \( \hat{V}_{Q, t+T} \) and \( \hat{V}_{Q, t+T}^2 \) at each observation time, for all the included tenors, based on the expressions in (21) and (22), respectively. Following the extant literature, we measure time in business days; i.e., the length of time from the end of trading on one day to end of trading on the following business day is set to 1/252. In addition, to account for the well-known intraday pattern in volatility, we compute time-to-maturity during a trading day on a “business time scale,” with the length of the within day windows set to equalize the contribution to the intraday volatility. To allow for possible dynamic changes in this intraday pattern, we further calculate this on a one-year rolling basis from the intraday and overnight realized volatilities over the past year. Using the regressions in (23), we then calculate the instantaneous variance estimates \( \hat{V}_t^Q \) and \( \hat{V}_t^Q \) as defined in (24). We rely on (26) for our computation of \( \hat{L}_t^Q \), with the \( \hat{s}_t \) adjustment term in (25) calculated from the high-frequency return and option observations on the specific trading day. For each trading day, we then average all the high-frequency estimates and henceforth rely on these as our daily estimates.

Our realized risk measures \( \hat{RV}_t \) and \( \hat{RL}_t \) are similarly computed over windows of one trading day. To allow for comparison with the risk-neutral measures, we further normalize
the high-frequency return measures using an overnight adjustment factor, based on the ratio of the intraday to the overnight realized volatilities computed on a one-year rolling basis. Table 1 reports the full-sample means for the resulting daily $\mathbb{P}$ and $\mathbb{Q}$ estimates, together with robust standard errors in parentheses.\footnote{We purposely report the 1\% truncated means here and throughout to avoid the full-sample values being unduly influenced by a few very large “crisis” observations.} Since all the distributions are heavily right skewed, we also report the corresponding quantiles.

Looking first at the volatility estimates in the top-portion of the table, the $\hat{V}_t^Q$ spot estimates are naturally lower than the short and long maturity $\hat{V}_{t,t+T}^Q$ estimates. Meanwhile, it is noteworthy that the relative difference between $\hat{V}_{t,t+T}^Q$ and $\hat{V}_t^Q$ is the smallest for the

<table>
<thead>
<tr>
<th>Statistic</th>
<th>Mean</th>
<th>$Q_{25}$</th>
<th>$Q_{50}$</th>
<th>$Q_{75}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\hat{V}_{t,t+T}^Q$, short $T$</td>
<td>0.0395</td>
<td>0.0130</td>
<td>0.0224</td>
<td>0.0429</td>
</tr>
<tr>
<td>$\hat{V}_{t,t+T}^Q$, long $T$</td>
<td>0.0418</td>
<td>0.0151</td>
<td>0.0248</td>
<td>0.0478</td>
</tr>
<tr>
<td>$\hat{V}_t^Q$</td>
<td>0.0378</td>
<td>0.0111</td>
<td>0.0201</td>
<td>0.0396</td>
</tr>
<tr>
<td>$\hat{R}V_t$</td>
<td>0.0282</td>
<td>0.0064</td>
<td>0.0127</td>
<td>0.0278</td>
</tr>
<tr>
<td>$3(\hat{V}<em>{t,t+T}^Q - \hat{V}</em>{t,t+T}^Q)$, short $T$</td>
<td>0.0028</td>
<td>0.0005</td>
<td>0.0010</td>
<td>0.0021</td>
</tr>
<tr>
<td>$3(\hat{V}<em>{t,t+T}^Q - \hat{V}</em>{t,t+T}^Q)$, long $T$</td>
<td>0.0060</td>
<td>0.0011</td>
<td>0.0019</td>
<td>0.0051</td>
</tr>
<tr>
<td>$3(\hat{V}_t^Q - \hat{V}_t^Q)$</td>
<td>0.0011</td>
<td>-0.0001</td>
<td>0.0002</td>
<td>0.0008</td>
</tr>
<tr>
<td>$-\hat{L}_t^Q$</td>
<td>0.0579</td>
<td>0.0089</td>
<td>0.0203</td>
<td>0.0464</td>
</tr>
<tr>
<td>$-\hat{R}L_t$</td>
<td>0.0331</td>
<td>0.0019</td>
<td>0.0056</td>
<td>0.0179</td>
</tr>
</tbody>
</table>

Note: The table reports truncated at 1\% sample means, with Newey-West robust standard errors in parentheses, and quantiles for the different daily risk estimates defined in Sections 4.1 and 4.2. The estimates are constructed from short-dated S&P 500 index options and high-frequency SPY returns spanning 2007-2020.
highest quantile. This may be explained by the fact that mean reversion in volatility and the leverage effect impact the term structure of $\hat{V}_{t,t+T}^Q$ in opposite directions, thus diminishing the relative difference between $\hat{V}_t^Q$ and $\hat{V}_{t,t+T}^Q$ when volatility is high. Also, even though there is obviously a large gap between $\hat{V}_t^Q$ and $\hat{V}_{t,t+T}^Q$, consistent with a positive $IVRP_t$ and the widely documented significant variance risk premium over longer horizons, the gap between $\hat{V}_t^Q$ and $\hat{V}_{t,t+T}^Q$ for short $T$ is also fairly large, pointing to a sizable risk premium for shocks to the instantaneous risk-neutral variance.

Turning next to the $3(\hat{V}_{t,t+T}^Q - \hat{V}_{t,t+T}^Q)$ estimates, and measures of asymmetry in the risk-neutral return distribution, reported in the mid-portion of the table, the values seemingly depend strongly on the horizon $T$. Recall from the discussion in Section 2, that the leading term in an asymptotic expansion of $3(V_{t,t+T}^Q - V_{t,t+T}^Q)$ for small $T$ equals $3 \int_\mathbb{R} (z^2 - 2(e^z - 1 - z)) \nu_t^Q(dz)$. Meanwhile, the second-order term in said expansion that depends on the horizon $T$ may be traced to mean-reversion in volatility and jump intensity, as well as the leverage effect. The pairwise differences in the sample means and the various quantiles of $3(\hat{V}_{t,t+T}^Q - \hat{V}_{t,t+T}^Q)$ for short and long tenors thus directly underscore the importance of these latter effects. Further supporting this conjecture, the mean and the median of the instantaneous $3(\hat{V}_t^Q - \hat{V}_t^Q)$ estimates are also substantially lower than those of $3(\hat{V}_{t,t+T}^Q - \hat{V}_{t,t+T}^Q)$ for short $T$. Looking at the $3(\hat{V}_t^Q - \hat{V}_t^Q)$ estimates, it is also worth noting that with $3(V_t^Q - V_t^Q)$ being approximately equal to the risk-neutral third moment of the jumps (recall equation (7)), the numbers in the table point to very little asymmetry in the implied $Q$ jump size distribution.

## 5.3 Leverage Risk and Pricing

The final two rows in Table 1 summarize the results for our leverage risk estimates. Consistent with the idea of a nontrivially sized instantaneous leverage risk premium $ILRP_t$, the sample mean of the risk-neutral $-\hat{L}_t^Q$ exceeds that of the realized $-\hat{R}L_t$ by more than twice of what the sample mean of the risk neutral spot variance $\hat{V}_t^Q$ exceeds that of the realized variance $\hat{R}V_t$. The different quantiles for $-\hat{L}_t^Q$ are also all substantially higher than the corresponding quantiles for $-\hat{R}L_t$.

To help visualize and further appreciate the pricing of the jump leverage effect, and how it differs from the pricing of variance risk, Figure 3 plots the time series of $\hat{V}_t^Q$ and $\hat{R}V_t$, 
while Figure 4 plots $-\hat{L}_t^Q$ and $-\hat{R}L_t$. All of the measures obviously varied quite significantly over the 2007-2020 sample period, generally increasing during periods of crisis. As is to be expected, the realized measures, $\hat{R}V_t$ and $-\hat{R}L_t$, also typically exceed their risk-neutral counterparts, $\hat{V}_t^Q$ and $-\hat{L}_t^Q$, at the onset of a crisis, with that ordering reversed in the aftermath of a crisis. As a case in point, during the financial crisis in the Fall of 2008, $\hat{R}V_t$ exceeded $\hat{V}_t^Q$ for the month of September and beginning of October, while $\hat{V}_t^Q$ exceeded $\hat{R}V_t$ by quite a wide margin for more than a year thereafter. This same general pattern is also evident for the $-\hat{L}_t^Q$ and $-\hat{R}L_t$ leverage risk measures.

Figure 3: The figure plots the ten-day moving averages of the daily $\hat{V}_t^Q$ and $\hat{R}V_t$ variance risk estimates.

Figure 4: The figure plots the ten-day moving averages of the daily $-\hat{L}_t^Q$ and $-\hat{R}L_t$ leverage risk estimates.

Meanwhile, the by far largest increase in all of the risk measures occurred in the Spring
of 2020 coincident with the start of the COVID-19 pandemic. At that time $\hat{V}_t^Q$ reached a peak of roughly doubled that of its peak observed during the financial crisis of 2008. The heightened values of $\hat{V}_t^Q$ in the Spring of 2020 was, however, noticeable shorter-lived than the highs attained during the 2008 financial crisis. While these same general features manifest in $-\hat{L}_t^Q$ as well, there are also some noticeable differences.\(^{22}\) Most obviously, $-\hat{L}_t^Q$ increased even more dramatically during the recent pandemic, reaching a high of almost five-fold its high observed during the 2008 financial crises. In other words, if anything it appears as if leverage risk has become an even bigger concern to investors more recently than it has been historically.

To more specifically illustrate the behavior of the volatility and leverage risks during periods of market turmoil, Figure 5 plots the SPY price and the instantaneous risk-neutral volatility measure $\hat{V}_t^Q$ at a five-minute frequency for the three weeks in the sample when the market fell by more than 10%, namely the week of November 5-12, 2008 (displayed in the first panel), the week of August 3-10, 2011 (displayed in the second panel), and the week of March 4-11, 2020 (displayed in the third panel).\(^{23}\) Figure 6 plots the SPY and the leverage risk measure $-\hat{L}_t^Q$ for the same three weeks. Looking first at Figure 5, there is obviously a strong negative correlation between the price changes and the changes in the $\hat{V}_t^Q$ spot volatility estimates during each of these three turbulent weeks. These large realized leverage effects also clearly manifest in the $-\hat{L}_t^Q$ leverage risk estimates displayed in Figure 6.

Although both of the two risk measures generally appear to be strongly negatively correlated with the price, there are also some important differences between the three episodes. In particular, while the average values of $\hat{V}_t^Q$ for the first week of November 2008 and the second week of March 2020 were very similar, at 64.0% and 61.2%, respectively, the volatility stayed within a relatively narrow range of around 55% - 70% for the first of the two weeks, while it fluctuated much more widely from a low of around 40% to a high of around 85% for the second of the two weeks. The volatility mean reversion and the leverage effect also

\(^{22}\)Incidently, these differences in the time series behavior of the variance and leverage risk measures also indicate that the jump risks have a much more complicated structure than portrayed by most parametric models hitherto employed in the literature, the popular double-jump stochastic volatility model discussed in Appendix C included.

\(^{23}\)To allow for easier interpretation, the volatility measures in Figure 5 are displayed in annualized volatility units, or $(252 \times \hat{V}_t^Q)^{1/2}$. The market also fell by more than 10% the first week of October 2008. However, due to issues with missing observations and quality of the options data, we omit that week from the display.
both manifested very differently for these two different episodes. For instance, while the average value of \((252 \times \hat{\sigma}_t^{Q})^{1/2}\), as proxied by the VIX, equalled 59.2\% for the first week of November 2008, it was “only” equal to 43.1\% for the second of March 2020. Recalling again the short-time expansions for the risk-neutral volatility measures in Section 2, this noticeably lower average value of the VIX for the second of the two weeks may naturally be attributed to stronger mean reversion during that week. As previously noted, this strong volatility mean reversion was also accompanied by more pronounced leverage effects during
the pandemic. Indeed, as Figure 6 shows, the estimates for $\hat{L}_t^Q$ are an order of magnitude larger for the second week of March 2020 compared to the first week of November 2008.

The above discussed evidence for the presence and pricing of jump leverage risk holds that shocks in the form of jumps to the stochastic volatility and/or the jump intensity should be priced. If so, we should also expect the mean of the instantaneous drift $m_Q^t(V_Q^t)$ to be positive. Consistent with that idea and the expansion in (9), the full-sample mean of the estimates for $\beta_{1,1}$ from the regression in (23) equals 0.0636, corresponding to an estimate of 0.1272 for $m_Q^t(V_Q^t)$. We can similarly estimate the forward quantity, say $E_Q^t(m_{t+T}^Q(V_Q^t))$ for some $T > 0$, by employing options recorded at time $t$ expiring shortly after time $t + T$. Doing so with options expiring soon after two weeks, the full-sample mean estimate of $E_Q^t(m_{t+T}^Q(V_Q^t))$ only equals 0.0050, compared to 0.1272 for the mean instantaneous drift. A formal test for the null hypothesis that the mean of $m_Q^t(V_Q^t)$ equals that of $E_Q^t(m_{t+T}^Q(V_Q^t))$ is also strongly rejected with a $t$-statistic of 4.34. This significantly lower estimate for $E_Q^t(m_{t+T}^Q(V_Q^t))$ compared to the estimate for $m_Q^t(V_Q^t)$ thus again suggests that the risk premium associated with the variation in $V_Q^t$ is mostly due to the compensation for short-lived spikes, or jumps.

To further buttress this thesis, as an additional independent robustness check we next look at the pricing of VIX index options.

6 Jump Leverage Risk Premiums Embedded in VIX Options

If the jump leverage effect constitutes an important source of risk that requires compensation by investors, then some of the jumps to volatility, namely the ones directly related to price jumps, should also be priced. In fact, based on existing empirical evidence (see, e.g., Jacod and Todorov (2010), Todorov and Tauchen (2011) and Caporin et al. (2017), among others), one would naturally expect most volatility jumps to be accompanied by simultaneous price jumps in the opposite direction.\footnote{This strong negative dependence between jumps in the market and the volatility of the market also naturally arises in many equilibrium models, the recent model by Eraker and Yang (2022) included. Counter to the empirical evidence, however, this particular theoretical model, as well as other exponentially-affine models, cannot account for negative volatility jumps; see also the discussion in Amengual and Xiu (2018).} As such, this allows for independent verification of the
importance of the jump leverage risk premium through the use of options written on the VIX index and the pricing thereof.

The VIX index computed by the CBOE is formally based on the \( \hat{V}^{Q}_{t,t+T} \) risk-neutral variance measure defined in (4) together with the two available tenors \( T \) closest to 30 calendar days.\(^{25}\) We obtain end-of-day prices for VIX options spanning the 2007-2020 period from OptionMetrics. We further collect 5-minute price records for the VIX index itself from the CBOE Data Shop. We apply standard cleaning procedures and similar filters to the ones discussed in Section 5.1 used for processing the S&P 500 price and options data.\(^{26}\) However, since the number of available short-tenor VIX index options is significantly less than the number of short-dated S&P 500 options, especially for the early part of the sample, we retain a somewhat wider set of VIX index options with maturities ranging from three up to thirty business days.

Using the VIX index option data and the expressions in (21) and (22), we begin by computing the two separate estimates for the risk-neutral variances of the VIX. We then run the regressions in (23), omitting the \( T^2 \) terms due to the more limited availability of different tenors. We refer to the resulting estimated intercepts, and instantaneous risk-neutral variance estimates of the log(VIX) index, by \( \hat{V}^Q_t \) and \( \hat{V}^Q_t \), respectively. In parallel to our analysis of the S&P 500 discussed above, we then compare these option-implied volatilities for the VIX with their daily realized counterparts,

\[
\hat{RVV}_t = \frac{n}{k_n} \sum_{i=1}^{k_n} (\Delta^{n}_{i,t} VIX)^2, \quad \hat{RVV}_t = \frac{2n}{k_n} \sum_{i=1}^{k_n} (e^{\Delta^{n}_{i,t} VIX} - 1 - \Delta^{n}_{i,t} VIX),
\]

where \( \Delta^{n}_{i,t} VIX \equiv \log(VIX_{t-i/n}) - \log(VIX_{t-(i-1)/n}) \) denote the intraday high-frequency logarithmic returns on the VIX.

Given the aforementioned strong empirical evidence for negative dependence between

\(^{25}\)The CBOE uses a calendar day convention in their computation of the VIX, while as previously noted we rely on a business day convention in our high-frequency calculations of \( \hat{V}^{Q}_{t,t+T} \). However, since we are not actually modeling the connection between the VIX and our \( \hat{V}^{Q}_{t,t+T} \) measure, this difference is immaterial.

\(^{26}\)We also further apply bounce-back filters similar to the ones proposed by Barndorff-Nielsen et al. (2009) to remove rare large intraday price swings caused in part by the rules applied by the exchange for the inclusion of deep out-of-the-money options in the computation of the VIX; see Andersen et al. (2015) for a more detailed discussion of this issue.
market price jumps and jumps in the VIX, together with the strong empirical evidence for the significant pricing of market price jumps, one would naturally expect volatility jump risk to be priced with the opposite sign to that of market price jumps. In particular, $3(\hat{VV}_Q - \hat{V}_Q)$, which by the discussion in Section 2 provides a proxy for the third risk-neutral moment of the jumps in the log-VIX index, would naturally be expected to be less than the corresponding realized measure $3(\hat{RVV}_t - \hat{RVV}_t)$, with the gap stemming from the pricing of volatility jump risk. Consistent with that conjecture, the full-sample mean of $3(\hat{VV}_Q - \hat{V}_Q)$ equals $-0.1102$, while that of $3(\hat{RVV}_t - \hat{RVV}_t)$ only equals $-0.0007$. The $t$-statistic of 19.16 for testing whether the risk-neutral and realized third moments of the jumps to the market volatility are the same is also highly significant. In other words, jumps to the state variables that drive the stochastic volatility and risk-neutral jump intensity of the market are obviously priced by investors, further corroborating the practical importance of the jump leverage effect and the sizeable magnitude of the corresponding jump leverage risk premium.

7 Concluding Remarks

Jumps in asset prices trigger two types of risks: (i) risk stemming from changes in the instantaneous volatility, and (ii) risk stemming from changes in the future investment opportunity set due to jumps in the diffusive volatility and/or jump intensity. We provide new model-free evidence in support of the nontrivial pricing of the latter type of jump risk, and so-called jump leverage risk in particular. Our analyses rely on novel short-time risk-neutral variance expansions, together with short-dated S&P 500 index options and high-frequency S&P 500 returns for estimating different risk measures and quantifying the corresponding risk premiums.

Our new model-free risk premium estimates also provide useful guidance for the future development of empirically more realistic equilibrium-based asset pricing models able to explain the significant pricing of jump risks. Most modern general equilibrium asset pricing models hold that the market and its instantaneous risk-neutral volatility are functions of some underlying state vector $S_t$:

$$X_t = F(S_t), \quad \mathcal{V}_t^Q = G(S_t).$$
The state vector $S_t$ is typically assumed to follow a Markov jump-diffusion. The specific variables entering $S_t$, and the functional forms of $F$ and $G$, will depend on the particular equilibrium model. The market price and its volatility, $X_t$ and $V^Q_t$, may also depend on different components of $S_t$. The exact form of these dependencies in turn determine the risks of the state variables that command risk premiums and the signs and connection between the instantaneous variance and leverage risk premiums, $\text{IVRP}_t$ and $\text{ILRP}_t$, emphasized here. To further clarify, it is instructive consider three different classes of equilibrium models that have recently been proposed in the literature to help explain option data and the pricing of jump risks.

- A rare disaster risk model with time-varying probabilities of disasters and a representative agent with Epstein-Zin utility (e.g., Wachter (2013) and Seo and Wachter (2019)). In this class of models, the aggregate consumption exhibits jumps (or rare disasters), which carries over to the aggregate dividend of the market. The state vector $S_t$ includes the aggregate dividend and the time-varying probability of jumps (or rare disasters). Importantly, $X_t$ depends on both state variables, while the risk-neutral spot variance $V^Q_t$ only depends on the second state variable. Consequently, even though $X_t$ is subject to occasional jumps, $X_t$ and $V^Q_t$ do not co-jump. Consequently, while $\text{IVRP}_t > 0$, in this class of models $\text{ILRP}_t \equiv 0$.

- A rare disaster risk model and a representative agent with external habit formation risk preferences (e.g., Du (2011)). In this model the aggregate consumption is again subject to jumps. The state variable $S_t$ is comprised of the aggregate level of consumption and the level of habit. Since the time-varying risk aversion of the representative agent is determined by habit formation, which depends on the value of the aggregate consumption, jumps in risk aversion occur at the same time as jumps in the aggregate consumption. A negative jump in consumption therefore triggers a positive jump in risk-aversion, which in turn implies a positive jump in $V^Q_t$. As a result, $\text{IVRP}_t > 0$ and $\text{ILRP}_t > 0$, both due to jumps in consumption.

- A model with jumps in the volatility of consumption growth and a representative agent with Epstein-Zin utility (e.g., Eraker and Shaliastovich (2008), Drechsler and Yaron
(2011) and Eraker and Yang (2022)). In this class of models, the state variable $S_t$ includes consumption growth, which does not jump, and the volatility of consumption growth, which varies over time and may jump. Since investors have Epstein-Zin preferences, shocks to the time-varying volatility, including volatility jumps, will be priced. Positive jumps in consumption volatility will therefore also trigger negative jumps in the market, as well as positive jumps in the risk-neutral variance $\nu_t^Q$. Consequently, $IVRP_t > 0$ and $ILRP_t > 0$, both due to jumps in consumption growth volatility.

Although formally distinct, the economic channels through which the jump leverage risk premium arise in the latter two equilibrium classes of models discussed above both depend on price jumps being accompanied by sudden changes in the investment opportunity set and/or the investor’s risk aversion. It would be interesting to further calibrate the models and quantify the magnitudes of these different channels to help precise the economic mechanisms that drive the empirically “large” jump leverage risk premium document here. We leave further work along these lines for future research.
Appendix: Derivations of Short-Term Variance Expansions

A.1 Assumptions

We set the risk-free rate and the dividend yield to zero and make the normalization $x_0 = 0$. We further denote the martingale component of $x$ with

$$ M_T = \int_0^T \sigma_s dW_s^Q + \int_0^T \int_\mathbb{R} z(\mu - \nu^Q)(ds,dz). \quad (A.1) $$

To derive the expansions, we assume that the $Q$-dynamics of $z$, for $z$ being one of the processes $V^Q, V^Q, m^Q(V^Q), m^Q(m^Q(V^Q)), L^Q, m^Q(L^Q), m^Q(V^Q)M, m^Q(m^Q(V^Q)M), [V^Q, V^Q]$ and $m^Q([V^Q, V^Q])$, is of the form

$$ z_t = z_0 + \int_0^t m^Q_s(z)ds + Q - \text{martingale}, \quad (A.2) $$

with the above martingale being $\mathcal{F}_0$-conditionally square-integrable, $m^Q_t$ being a process with càdlàg paths, and $\mathbb{E}^Q_0(m^Q_t)^2 < \infty$. We further denote the martingale part of $V^Q$ by $Z^Q(V^Q)$. We assume that the above decomposition also holds for the processes $m^Q(V^Q)Z^Q(V^Q)$, $m^Q(m^Q(V^Q))Z^Q(V^Q)$ and $m^Q(m^Q(V^Q))Z^Q(V^Q)$.

The above assumptions are rather weak and generally satisfied for the continuous-time asset pricing models used in finance, including standard reduced-form models, as well as equilibrium-based models. In particular, the assumptions hold when the asset price dynamics is embedded in a general stochastic differential equation (SDE), and the coefficients of the SDE have finite conditional moments of certain order.
A.2 Proofs

Given the dynamics of $x$ under $Q$ and using the notation of $M$ above, it follows by an application of Itô’s formula that

$$x_T = -\frac{1}{2} \int_0^T V_s^Q ds + M_T. \tag{A.3}$$

Another application of Itô’s formula for the product of two processes, leads to

$$T \times V_{0,T}^Q = \mathbb{E}_0^Q \left( -\frac{1}{2} \int_0^T (V_s^Q - \mathbb{E}_0^Q(V_s^Q)) ds + M_T \right)^2$$

$$= \frac{1}{4} \text{var}_0^Q \left( \int_0^T V_s^Q ds \right) + \mathbb{E}_0^Q (M_T^2) - \mathbb{E}_0^Q \left( \int_0^T V_s^Q ds M_T \right)$$

$$= \frac{1}{4} \text{var}_0^Q \left( \int_0^T V_s^Q ds \right) + \int_0^T \mathbb{E}_0^Q(V_s^Q) ds - \int_0^T \text{cov}_0^Q(V_s^Q, M_s) ds, \tag{A.4}$$

and

$$T \times V_{0,T}^Q = \int_0^T \mathbb{E}_0^Q(V_s^Q) ds. \tag{A.5}$$

Using our assumption for $V_s^Q$, $m^Q(V^Q)$ and $m^Q(m^Q(V^Q))$, we further have

$$\mathbb{E}_0^Q(V_s^Q) = \int_0^s \mathbb{E}_0^Q(m_u^Q(V^Q)) du = \int_0^s \mathbb{E}_0^Q \left( m_0^Q(V^Q) + \int_0^u m_v^Q(m^Q(V^Q)) dv \right) du$$

$$= \int_0^s \mathbb{E}_0^Q \left( m_0^Q(V^Q) + u \times m_0^Q(m^Q(V^Q)) + \int_0^u \int_0^v m_z^Q(m^Q(m^Q(V^Q))) dz dv \right) du$$

$$= s \times m_0^Q(V^Q) + s^2 \times C_0 + \tilde{C}_0(s), \tag{A.6}$$

for some $\mathcal{F}_0$-adapted random variable $C_0$ and $\mathcal{F}_0$-adapted random function $\tilde{C}_0(s)$ satisfying $|\tilde{C}_0(s)| \leq s^3 \times \tilde{C}_0''$, where $\tilde{C}_0''$ is another $\mathcal{F}_0$-adapted random variable. A similar expansion for $\mathbb{E}_0^Q(V_s^Q)$ also readily obtains. Next, applying Itô’s formula and using the definition of the processes $M$ and $L^Q$ (note in particular that $M_0 = 0$), as well as our assumptions for the
dynamics of $L^Q, m^Q(L^Q), m^Q(V^Q)M$ and $m^Q(m^Q(V^Q)M)$, we obtain

$$
\text{cov}^Q_0(\mathcal{V}_s, M_s) = \mathbb{E}^Q_0 \left( \int_0^s m_u^Q(\mathcal{V}^Q)duM_s \right) + \mathbb{E}^Q_0(\mathcal{Z}_s^Q(\mathcal{V}^Q)M_s) \\
= \mathbb{E}^Q_0 \left( \int_0^s m_u^Q(\mathcal{V}^Q)M_u \right) + \mathbb{E}^Q_0 \left( \int_0^s \mathcal{L}_u^Q \right) \\
= \mathbb{E}^Q_0 \left( \int_0^s \int_0^u m_v^Q(m^Q(V^Q)M)dvdu \right) + \mathbb{E}^Q_0 \left( \int_0^s \left( \mathcal{L}_u^Q + \int_0^u m_v^Q(L^Q)dv \right) du \right) \\
= s \times L_0^Q + s^2 \times C_0 + \tilde{C}_0(s),
$$

(A.7)

where $C_0$ and $\tilde{C}_0(s)$ satisfy the same properties as above. Finally, direct expansion and the fact that $Z^Q(V^Q)$ is a martingale leads to:

$$
\text{var}^Q_0 \left( \int_0^T \mathcal{V}_s^Q ds \right) = \mathbb{E}^Q_0 \left( \int_0^T \int_0^s \left( m_u^Q(\mathcal{V}^Q) - \mathbb{E}^Q_0(m_u^Q(\mathcal{V}^Q)) \right) duds \right)^2 \\
+ \mathbb{E}_0 \left( \int_0^T \int_0^T Z_u^Q(\mathcal{V}^Q)Z_v^Q(\mathcal{V}^Q)dudv \right) \\
+ 2\mathbb{E}^Q_0 \left( \int_0^T \int_0^s \left( m_u^Q(\mathcal{V}^Q) - \mathbb{E}^Q_0(m_u^Q(\mathcal{V}^Q)) \right) duds \int_0^T Z_u^Q(\mathcal{V}^Q))du \right).
$$

(A.8)

Using the definition of quadratic variation and the fact that $Z^Q(V^Q)$ is a martingale, we may write:

$$
\mathbb{E}_0 \left( \int_0^T \int_0^T Z_u^Q(\mathcal{V}^Q)Z_v^Q(\mathcal{V}^Q)dudv \right) = 2 \int_0^T (T - s)\mathbb{E}^Q_0([\mathcal{V}^Q,\mathcal{V}^Q]_s)ds.
$$
Next, using again the fact that $Z^Q(\mathcal{V}^Q)$ is a martingale,

$$
\mathbb{E}_0^Q \left( \int_0^T \int_0^s (m_u^Q(\mathcal{V}^Q) - \mathbb{E}_0^Q(m_u^Q(\mathcal{V}^Q))) \, du \int_u^T Z_u^Q(\mathcal{V}^Q) \, duds \right)
$$

$$
= \mathbb{E}_0^Q \left( \int_0^T (T - s) \int_0^s (m_u^Q(\mathcal{V}^Q) - \mathbb{E}_0^Q(m_u^Q(\mathcal{V}^Q))) \, du \int_u^s Z_u^Q(\mathcal{V}^Q) \, duds \right)
$$

$$
+ \mathbb{E}_0^Q \left( \int_0^T \int_0^s (m_u^Q(\mathcal{V}^Q) - \mathbb{E}_0^Q(m_u^Q(\mathcal{V}^Q))) \, du \int_0^s Z_u^Q(\mathcal{V}^Q) \, duds \right)
$$

$$
= 2\mathbb{E}_0^Q \left( \int_0^T (T - s)m_s^Q(\mathcal{V}^Q)Z_s^Q(\mathcal{V}^Q) \right) \, ds + \mathbb{E}_0^Q \left( \int_0^T \int_0^s (m_s^Q(\mathcal{V}^Q) - m_u^Q(\mathcal{V}^Q))Z_u^Q(\mathcal{V}^Q) \, duds \right).
$$

From here, by making use of our assumptions for $[\mathcal{V}^Q, \mathcal{V}^Q]$, $m^Q([\mathcal{V}^Q, \mathcal{V}^Q])$, $m^Q(\mathcal{V}^Q)Z^Q(\mathcal{V}^Q)$, $m^Q(m^Q(\mathcal{V}^Q))$, $m^Q(m^Q(\mathcal{V}^Q))Z^Q(\mathcal{V}^Q)$ and $m^Q(m^Q(\mathcal{V}^Q))Z^Q(\mathcal{V}^Q)$, it now follows that

$$
\text{var}_0^Q \left( \int_0^T \mathcal{V}_s^Q \, ds \right) = T^3 \times C_0 + \tilde{C}_0(T),
$$

where $C_0$ is some $\mathcal{F}_0$-adapted random variable, and the $\mathcal{F}_0$-adapted random function $\tilde{C}_0(T)$ satisfies $|\tilde{C}_0(T)| \leq T^4 \times \tilde{C}_0'$ for some some $\mathcal{F}_0$-adapted random variable $\tilde{C}_0'$.

**B Appendix: Nonzero Interest Rate and Dividend Yield**

All of our expansions and empirical calculations are based on the simplifying assumption that the risk-free rate and the dividend yield are identically equal to zero. In this appendix, we show how the empirical results are practically unaffected by this assumption.

We denote the constant continuously-compounded risk-free rate and dividend yield by $r$ and $d$, respectively. We further denote the time $t$ price of a futures contract written on the market expiring at time $t + T$ by $F_{t,T}$, with the corresponding log price denoted by $f_{t,T}$. By standard no-arbitrage pricing arguments, $F_{t,T} = e^{(r-d)T}X_t$. With this additional notation, the definitions of our two risk-neutral variance measures previously defined in (3) and (4)
now naturally become

\[ V_{t,t+T}^Q = \frac{1}{T} \text{Var}_t^Q(x_{t+T} - f_{t,T}), \quad \gamma_{t,t+T}^Q = -\frac{2}{T} \mathbb{E}_t^Q(x_{t+T} - f_{t,T}). \]  \hspace{1cm} (B.1)  

Note that since \( f_{t,T} \) is adapted to \( \mathcal{F}_t \), the definition of \( V_{t,t+T}^Q \) remains unchanged compared to the case with \( r = d = 0 \), while the centering term in \( \gamma_{t,t+T}^Q \) gets slightly modified. With this adjustment, the key expansion results in (9) and (13) both continue to hold.

Correspondingly, the option spanning results needed to construct the estimates of \( V_{t,t+T}^Q \) and \( \gamma_{t,t+T}^Q \) now take the form

\[ \mathbb{E}_t^Q(x_{t+T} - f_{t,T})^2 = 2 \int_0^\infty \left( 1 - \log \left( \frac{K}{X_t} \right) \right) \frac{e^{rT} O_{t,T}(K)}{K^2} dK, \]  \hspace{1cm} (B.2)

and

\[ \mathbb{E}_t^Q(x_{t+T} - f_{t,T}) = -\int_0^\infty \frac{e^{rT} O_{t,T}(K)}{K^2} dK. \]  \hspace{1cm} (B.3)

Compared to the previous expressions in (19) and (20), the only difference is that we now consider the future values of the option prices. Using these modified spanning results, \( \hat{V}_{t,t+T}^Q \) and \( \hat{\gamma}_{t,t+T}^Q \) may again be constructed exactly as previously done in (21) and (22).

Importantly, the above analysis further implies that if nonzero interest rates and dividend yields are ignored, then, up to a higher order term, the OLS estimates of \( b_{t,0}, b_{t,1}, \beta_{t,0} \) and \( \beta_{t,1} \) from the regressions in (23) will simply be scaled down by the same interest rate factor \( e^{-rT} \) relative to their counterparts where nonzero \( r \) and \( d \) are explicitly accounted for. Obviously, 

\[ e^{-rT} = 1 - rT + O(T^2) \]  as \( T \downarrow 0 \). Hence, not only will this correction similarly affect all the coefficient estimates of interest, given the invariably low value of the risk-free interest rate, it will also be numerically very small.
C Appendix: Jumps and Leverage Risk in a Parametric Model

This Appendix demonstrates how the risk premium measures defined in Section 3 manifest in the popular parametric double-jump stochastic volatility model of Duffie et al. (2000). This particular model has been extensively used in the empirical asset pricing literature.

The model postulates the following asset dynamics under \( \mathbb{P} \)

\[
\frac{dX_t}{X_{t-}} = \alpha_t dt + \sigma_t dW^P_t + \int_{\mathbb{R}^2} (e^z - 1)(\mu - \nu^P)(ds, dz, dy),
\]

\[
d\sigma^2_t = \kappa^P(\theta^P - \sigma^2_t)dt + \eta \sigma_t dB^P_t + \int_{\mathbb{R}^2} y(\mu - \nu^P)(ds, dz, dy),
\]

where \( \alpha_t \) is a linear function of \( \sigma^2_t \), \((W^P_t, B^P_t)\) is a bivariate Brownian motion with correlation \( \rho \), and the jump compensator takes the form

\[
\nu^P(ds, dz, dy) = (\lambda_0^P + \lambda_1^P \sigma^2_t) e^{-\frac{(z-\mu_z^P)^2}{2\nu^2_z}} e^{-\frac{y}{\mu^P_y}} \frac{1_{\{y>0\}}}{\sqrt{2\pi \nu^P_z \mu^P_y}} ds dz dy.
\]

This specification in turn accounts for a number of key empirically relevant features: volatility is time-varying, price and volatility can both jump, and the jumps typically arrive together, the intensity of the jumps is an affine function of \( \sigma^2_t \) and therefore time-varying. We note that we have slightly constrained the original specification of Duffie et al. (2000) by assuming that the size of the price and volatility jumps are independent of each other, even though they arrive together. This simplification of the double-jump volatility model is only for ease of exposition and has no bearing on our analysis of jump leverage.

Per the discussion in the main text, assume for simplicity that the risk-free interest rate and the dividend yield are both zero. The dynamics of \( X \) under the \( \mathbb{P} \) and \( \mathbb{Q} \) measures then formally coincide, although the pricing of various risks may cause the values of the risk-neutral parameters, superscripted with \( \mathbb{Q} \) in the following, to differ from the corresponding \( \mathbb{P} \) superscripted parameters. Importantly, however, the \( \rho \) and \( \eta \) parameters do not change under this equivalent change of measure.
To begin, consider the instantaneous variance risk premium. Using the shorthand notation \( \tilde{v}_z^Q = 2 \left( e^{\mu^Q_0 + v^Q_z/2} - 1 - \mu^Q_0 \right) \), the two differently defined risk-neutral instantaneous variances may be expressed as

\[
V^Q_t = \sigma^2_t + (\lambda^Q_0 + \lambda^Q_1 \sigma^2_t)((\mu^Q_z)^2 + \tilde{v}^Q_z), \quad \mathcal{V}^Q_t = \sigma^2_t + (\lambda^Q_0 + \lambda^Q_1 \sigma^2_t)\tilde{v}^Q_z.
\] (C.4)

The same expressions also obtain for \( V^P_t \) and \( \mathcal{V}^P_t \), except for the \( Q \)-parameters being replaced by their \( P \) counterparts. Correspondingly, the \( IVR P_t \) instantaneous variance risk premium takes the specific form

\[
IVR P_t = (\lambda^Q_0 + \lambda^Q_1 \sigma^2_t)((\mu^Q_z)^2 + \tilde{v}^Q_z) - (\lambda^P_0 + \lambda^P_1 \sigma^2_t)((\mu^P_z)^2 + \tilde{v}^P_z).
\] (C.5)

As previously noted, \( IVR P_t \) generally only reflects the pricing of price jump risk, which in the context of the double-jump stochastic volatility model can arise from either different jump intensities, or different jump distributions under the \( P \) and \( Q \) measures.

Turning next to the leverage risk, we have

\[
\frac{1}{T} \left( [x, V^Q]_{t+T} - [x, V^Q]_t \right) = \frac{1}{T} (1 + \lambda^Q_1 \tilde{v}^Q_z) \left( \rho_{\eta} \int_t^{t+T} \sigma^2_s ds + \sum_{s \in [t, t+T]} \Delta x_s \Delta \sigma^2_s \right).
\] (C.6)

The dependence of the leverage risk on the \( Q \)-parameters stems from the fact that this risk is defined as a co-movement between \( x \) and \( V^Q \), which depends on the \( Q \)-parameters because of jumps in the price and the pricing thereof. Meanwhile, the first term in the parentheses above stems from the dependence between the Brownian motions driving the price and volatility, while the second term is due to the price-volatility co-jumps. Taking conditional expectations and letting \( T \downarrow 0 \), it further follows that

\[
L^Q_t = (1 + \lambda^Q_1 \tilde{v}^Q_z) \left[ \rho_{\eta} \sigma^2_t + \mu^Q_x \mu^Q_y (\lambda^Q_0 + \lambda^Q_1 \sigma^2_t) \right],
\] (C.7)

and

\[
L^P_t = (1 + \lambda^P_1 \tilde{v}^P_z) \left[ \rho_{\eta} \sigma^2_t + \mu^P_x \mu^P_y (\lambda^P_0 + \lambda^P_1 \sigma^2_t) \right].
\] (C.8)

These expressions for the instantaneous leverage effects in turn implies the following instan-
taneous leverage risk premium

\[
ILRP_t = (1 + \lambda_1^Q \tilde{v}_z^Q) \left[ \mu_P^P \mu_Q^Q (\lambda_0^Q + \lambda_1^Q \sigma_t^2) - \mu_Q^Q \mu_P^P (\lambda_0^P + \lambda_1^P \sigma_t^2) \right]. \tag{C.9}
\]

Since the terms in \( L_t^Q \) and \( L_t^P \) due to the continuous leverage effect coincide, these terms cancel in the \( ILRP_t \) risk premium, which therefore only depends on price-volatility co-jumps. As such, \( ILRP_t \) is effectively a difference of, re-scaled by constants, \( P \) and \( Q \) jump intensities. This also directly parallels the expression for \( IVRP_t \). Correspondingly, \( IVRP_t \) and \( ILRP_t \) are both affine functions of \( \sigma_t^2 \), and thus exhibit the same dynamics. This result is a particular feature of the double-jump volatility model. If the intensity of the price jumps that arrive together with the volatility jumps differed from those that do not, this would naturally cause the dynamics of \( IVRP_t \) and \( ILRP_t \) to differ.

Finally, it is easy to see that

\[
m_P^P (\mathcal{V}_t^Q) = (1 + \lambda_1^Q \tilde{v}_z^Q) \kappa_P^P (\theta_P^P - \sigma_t^2), \quad m_Q^Q (\mathcal{V}_t^Q) = (1 + \lambda_1^Q \tilde{v}_z^Q) \kappa_Q^Q (\theta_Q^Q - \sigma_t^2), \tag{C.10}
\]

which implies the following change in the instantaneous drift term of \( \mathcal{V}_t^Q \)

\[
m_t^Q (\mathcal{V}_t^Q) - m_t^P (\mathcal{V}_t^Q) = (1 + \lambda_1^Q \tilde{v}_z^Q) [\kappa_Q^Q (\theta_Q^Q - \sigma_t^2) - \kappa_P^P (\theta_P^P - \sigma_t^2)]. \tag{C.11}
\]

This in turn implies the following compensation for jumps risk

\[
(1 + \lambda_1^Q \tilde{v}_z^Q) \int_{\mathbb{R}^2} y (\nu_t^Q (dz, dy) - \nu_t^P (dz, dy)) = (1 + \lambda_1^Q \tilde{v}_z^Q) \left[ \mu_Q^Q (\lambda_0^Q + \lambda_1^Q \sigma_t^2) - \mu_P^P (\lambda_0^P + \lambda_1^P \sigma_t^2) \right]. \tag{C.12}
\]

This expression, not surprisingly, looks very similar to the expression for \( ILRP_t \) in (C.9) above, as they both reflect the compensation for the simultaneously arriving jumps in price and volatility.

To more concretely illustrate the important role played by price-volatility co-jumps in the pricing of jump risk, we finish our discussion of the parametric double-jump stochastic volatility model by showing the model-implied term structure of the second and third risk-neutral moments with and without volatility jumps. To help more clearly illustrate the effect, we further plot the different term structures obtained for initial low and high levels.
of volatility, as defined relative to the unconditional model-implied volatility. We rely on the same parameter values listed in Table D.1 that we use in our Monte Carlo simulations discussed Appendix D.

Looking first at the left panels in Figures C.1 and C.2 with and without volatility jumps, respectively, the general features of the term structures for the second moments appear largely unaffected by volatility jumps. By comparison, the term structures of the third risk-neutral moments clearly differ across the two figures, and are obviously much steeper when volatility jumps are allowed for. In fact, the slope of the term structure appears almost flat in the high volatility regime in the absence of volatility jumps, while the slopes are highly significant and positive regardless of whether the initial volatility is low or high when volatility jumps are included. Importantly, the patterns evident in Figure C.1 also fairly closely mirror the corresponding actual empirical illustration in Figure 2, thus indirectly underscoring the importance of price-volatility co-jumps, and the jump leverage effect, from a practical pricing perspective.
Figure C.1: The figure plots the model-implied term structure of the risk-neutral second and third moments allowing for volatility jumps. The risk-neutral moments are computed from the double-jump stochastic volatility model with the parameters given in Table D.1. The low and high volatility regimes corresponds to $V_0 = 0.017$ and $V_0 = 0.035$, respectively. The moments at the shortest horizons in each of the panels are normalized to unity.
Figure C.2: The figure plots the model-implied term structure of the risk-neutral second and third moments not allowing for volatility jumps. The risk-neutral moments are computed from the double-jump stochastic volatility model with parameters given in Table D.1, except for $\mu^Q_y$ which is fixed at zero. The low and high volatility regimes correspond to $V_0 = 0.017$ and $V_0 = 0.035$, respectively. The moments at the shortest horizons in each of the panels are normalized to unity.
D Appendix: Monte Carlo Simulations

This appendix evaluates the accuracy of the new $\hat{\nu}_t^Q$, $\hat{\rho}_t^Q$ and $\hat{\lambda}_t^Q$ feasible risk-neutral measures discussed in Section 4.1. We rely on the double-jump stochastic volatility model discussed in Appendix C to generate the true option prices. The specific values of the model parameters used in the simulations are reported in Table D.1.

Table D.1: Parameter Values

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<tr>
<td>$\rho$</td>
<td>$-0.9$</td>
<td>$\lambda_0^Q$</td>
<td>0</td>
<td>$\lambda_1^Q$</td>
<td>385</td>
</tr>
<tr>
<td>$\mu_z^Q$</td>
<td>$-0.05$</td>
<td>$\mu_y^Q$</td>
<td>0.0234</td>
<td>$\sqrt{v_z^Q}$</td>
<td>0.01</td>
</tr>
</tbody>
</table>

*Note:* The table reports risk-neutral parameter values for the double-jump stochastic volatility model in (C.1)-(C.3) in Appendix C used in the simulations.

The parameter values imply that $E^Q(\sigma_t^2) = 0.0257$, together with a relatively fast mean reversion of volatility with a half-life of around 8 days only. The jump parameters are set in a way so that $\int_{R^2} z^2 \nu_t(dz,dy) = \sigma_t^2$, implying that jumps contribute half of the risk-neutral jump variation at any point in time. We also allow for nontrivial volatility jumps, with a mean value equal to 0.0234, very close to that of the risk-neutral mean of $\sigma_t^2$. The observed option prices used in the estimation are contaminated by the following errors

$$\hat{O}_{t,T}(K_j) = O_{t,T}(K_j)(1 + 0.025 \times z_{t,T}(j)), \quad j = 1, \ldots, N_{t,T},$$

where $\{z_{t,T}(j)\}_{j=1}^{N_{t,T}}$ are sequences of *i.i.d.* standard normal independent random variables. The size of the observation error is calibrated to roughly match the bid-ask spreads of the S&P 500 index options used in our empirical analyses.

We initialize the simulations by setting the time-$t$ value of the spot variance to a low, average and a high value of 0.0170, 0.0204 and 0.0267, corresponding to the 25-th, 50-th and 75-th quantiles, respectively, of the unconditional distribution of $\sigma_t^2$. The initial level of the underlying stock price is set to 4,500. For each $(t,T)$ pair the strike grid is equidistant.
with gaps between strikes of 5. Starting from the at-the-money strike of 4,500, the strikes are then extended on both sides until the true out-of-the-money option price falls below 0.075. We employ three different tenors in the estimation, namely $T_1 = 3/252$, $T_2 = 5/252$ and $T_3 = 10/252$, corresponding to 3, 5 and 10 business days to expiration, respectively. Again, this specification of the strike grids and the choice of tenors mimics those of the S&P 500 index options used in the actual estimation. Finally, the $\widehat{s}_t$ bias-adjustment term for $\widehat{L}_t^Q$ is computed from 80 high-frequency return and option observations, approximately corresponding to a day worth of observations when sampling at the 5-minute frequency.

Table D.2: Monte Carlo Simulation Results

<table>
<thead>
<tr>
<th>Estimand</th>
<th>True Value</th>
<th>Estimates</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Q_{25}</td>
<td>Q_{50}</td>
</tr>
<tr>
<td>Initial Volatility $\sigma_t^2 = 0.0170$</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$V_t^Q$</td>
<td>0.0340</td>
<td>0.0336</td>
</tr>
<tr>
<td>$Y_t^Q$</td>
<td>0.0337</td>
<td>0.0333</td>
</tr>
<tr>
<td>$-L_t^Q$</td>
<td>0.0214</td>
<td>0.0212</td>
</tr>
<tr>
<td>Initial Volatility $\sigma_t^2 = 0.0204$</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$V_t^Q$</td>
<td>0.0408</td>
<td>0.0403</td>
</tr>
<tr>
<td>$Y_t^Q$</td>
<td>0.0404</td>
<td>0.0400</td>
</tr>
<tr>
<td>$-L_t^Q$</td>
<td>0.0257</td>
<td>0.0245</td>
</tr>
<tr>
<td>Initial Volatility $\sigma_t^2 = 0.0267$</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$V_t^Q$</td>
<td>0.0534</td>
<td>0.0528</td>
</tr>
<tr>
<td>$Y_t^Q$</td>
<td>0.0529</td>
<td>0.0524</td>
</tr>
<tr>
<td>$-L_t^Q$</td>
<td>0.0336</td>
<td>0.0305</td>
</tr>
</tbody>
</table>

*Note:* The table reports the simulated quantiles of the $V_t^Q$, $Y_t^Q$, and $-L_t^Q$ risk estimates for different initial volatility levels based on 1,000 Monte Carlo replications.

Table D.2 reports the quantiles of the resulting $\widehat{V}_t^Q$, $\widehat{Y}_t^Q$ and $\widehat{L}_t^Q$ risk estimates based on 1,000 Monte Carlo replications. As the table shows, the new estimates generally perform admirably, with the medians almost exactly equal to the true values, together with fairly narrow interquartile ranges.
References


