State-Dependent Intertemporal Risk-Return Tradeoff: Further Evidence

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Abstract

We suggest that the intertemporal risk-return tradeoff is not necessarily positive but rather state dependent. We further explore the state dependent risk-return relation by examining how the positive risk-return relation is distorted in response to various market conditions, including extreme price changes, differing levels of investor sentiment, the introduction of stock options, and throughout business cycles. The tendency for uninformed investors to be optimistic (pessimistic) in response to good (bad) market news cause overpricing (underpricing), and the resulting trade activity of arbitrageurs that distorts the positive risk-return tradeoff, is documented consistently across these environments. We find that the attenuation (reinforcement) of the positive risk-return relation under investors' optimistic (pessimistic) expectations is stronger in high (low) sentiment periods, in the presence of extreme returns, in the period after stock options became available, and during expansionary periods. We argue that the asymmetric intertemporal risk-return relation is a consequence of rational arbitrageurs' trading to exploit mispricing through the selling of overpriced stocks conditional on good news and buying underpriced stocks conditional on bad news.

JEL classification: G10; G12

Keywords: Intertemporal risk-return relation; Optimistic and pessimistic expectation; Overpricing and underpricing; Investor sentiment; Business cycle; Short selling

1. Introduction

The tradeoff between risk and return is one of the core tenets of financial economics. In particular, the intertemporal risk-return relation is vital for the time-varying rational expectations hypothesis which implies that rational risk-averse investors revise their expectations in response to changing volatility. Despite its importance in asset pricing theories, the actual sign of the intertemporal risk-return relation has been debated for decades, with empirical findings that are mixed and inconclusive. While a positive risk-return tradeoff is consistent with theoretical predictions, some argue that in reality it can be close to zero or even negative.

Recent studies provide empirical evidence that the sign of the ex-ante mean-variance relation is state-dependent. Yu and Yuan (2011) find that the relation is strongly positive during periods of low market sentiment but is close to zero when market sentiment is high. To explain their result, they argue that sentiment-driven traders have a greater effect on prices in periods of high sentiment due to their reluctance to act on low sentiment through short positions, and that sentiment-driven traders are also more likely to be naïve and mis-forecast the conditional volatility of returns. The result is that sentiment-driven traders undermine what would otherwise be a positive risk-return relation when sentiment is high. Extending this research, Stambaugh, Yu, and Yuan (2015) show that the combined effects of arbitrage risk (i.e., risk that deters arbitrage) and arbitrage asymmetry (i.e., relatively less arbitrage activity directed towards overpriced versus underpriced stocks) can induce a negative relation between aggregate idiosyncratic volatility and expected market return. They show that the effect of idiosyncratic volatility on expected return is strongly negative for overpriced stocks, but positive for underpriced stocks. In aggregate, the negative risk-return relation among overpriced stocks dominates the positive risk-return relation for underpriced stocks, which generates an overall negative risk-return relation.[1] Additional

[1] They also show that high (low) sentiment strengthens the negative (positive) relation among overpriced (underpriced) stocks, thus inducing a weak or negative (positive) relation between idiosyncratic volatility and expected return during periods of high (low) market sentiment.

evidence of state-dependency in the intertemporal risk-return relation is provided by Marks and Nam (2018), who show that good (bad) market news attenuates (strengthens) the positive risk-return relation in the short-term such that the intertemporal risk-return relation is weak or negative (significantly positive). Stressing the importance of a price adjustment process to capture the tendency for investors to correct prior mispricing when modeling return dynamics, they also show that ignoring this price adjustment process in the estimation of the risk-return relation leads to model misspecification which causes a significant upward (downward) bias in estimates of the relative risk aversion parameter conditional on good (bad) market news.[2]

The documented patterns in the state-dependent risk-return relation, and in particular the distortion of the positive intertemporal risk-return relation, have important theoretical and practical implications in behavioral finance. It is thus worthwhile to further investigate the consistency of the patterns that indicate state-dependency, their relative strengths, and explanations for the sources of these patterns. In this paper, we extend the aforementioned studies to further explore whether state dependency in the risk-return relation is induced in response to various market conditions. We contribute to the literature by adding further evidence that the distortion affecting the risk-return relation is present in a variety of market conditions and environments, and that the patterns of state-dependency are consistent with the notion that uninformed investors react optimistically (pessimistically) to good (bad) news.

In particular, we conduct the following empirical tests. First, we examine the robustness of a state-dependent risk-return tradeoff using different data sets, different sample periods, and the presence of extreme price changes. Second, we examine whether there is asymmetry in market imperfections that induces investors' optimistic (pessimistic) expectations about the future performance of stocks under good (bad) market news. Third, we examine how the observed asymmetric risk-return relation under good and bad market news varies across high and low sentiment periods. Fourth, since there is a significant relationship between time-varying stock market volatility and fluctuations in the level of real economic

[2] Good (bad) market news refers to a prior increase (decrease) in daily excess market returns, as in Marks and Nam (2018).

activity, we examine whether the business cycle affects uninformed investors' optimistic and pessimistic expectations and associated potential distortion of the positive risk-return relation under good and bad market news. Lastly, considering that short selling and options trading can be used by informed traders to earn arbitrage profits, we examine if the magnitude of the distortion of the positive risk-return relation is affected by the introduction of stock options.

The empirical models for returns we study are based on the asymmetric intertemporal risk-return model specifications suggested by both Yu and Yuan (2011) and Marks and Nam (2018). These studies employ Merton's (1973) simple linear relation between the expected market risk premium and conditional market volatility as a base model for measuring the intertemporal risk-return relation.[3] While Yu and Yuan (2011) employ an ad-hoc specification of dummies to capture high and low sentiment periods, the model of Marks and Nam (2018) is based on the Slutsky equation derived from a utility maximization model in which portfolio choice and equilibrium asset demand are jointly determined by two pricing channels, the conventional intertemporal risk-return tradeoff and investor adjustment behavior to correct prior mispricing. In the present work, we use the asymmetric intertemporal regression model suggested by Marks and Nam (2018) as a base model and incorporate dummies to capture possible state-dependency caused by the various market conditions described above. We especially explore the following three points that are not discussed in Yu and Yuan (2011) and Marks and Nam (2018). First, we check if the *constant* RRA parameter is a biased estimate of the true risk-return relation under good and bad market news. Second, we examine whether the estimated asymmetric market imperfections are consistent with optimistic and pessimistic expectations under good and bad market news. Third, we test whether the results of asymmetric RRA parameters along with optimistic and pessimistic expectations support our

[3] Merton (1973) proposed that the expected market risk premium is a linear function of its conditional variance and its covariance with investment opportunities. Merton (1980) showed that when the hedge component related to time-varying investment opportunities is negligible, the conditional market risk premium is proportional to conditional market volatility.

conjecture that the distortion of the positive risk-return relation is attributed to arbitrageurs' short-selling (buying) of overpriced (underpriced) stocks conditional on good (bad) market news.

Our work demonstrates that distortion of the risk-relation is pervasive and follows patterns of state-dependency across market environments that coincide with what one would expect based on mispricing. First, we find strong evidence of uninformed investors' optimistic and pessimistic expectations in response to good and bad market news, respectively, implying that optimistic (pessimistic) expectations about future stock performance induce overpricing (underpricing). Thus, as arbitrageurs trade to exploit this mispricing, the intertemporal mean-variance relation declines conditional on optimistic market conditions but becomes increasingly positive conditional on pessimistic market conditions. This implies that a relative risk aversion (RRA) parameter that is constant, and not dependent on market news, is a biased estimate of the intertemporal risk-return relation. Relative to using a more general asymmetric nonlinear specification to measure the risk-return tradeoff, the constant intertemporal risk-return relation *overestimates* the RRA parameter conditional on good market news, while *underestimating* the RRA parameter conditional on bad market news. Third, while good market news in high-sentiment periods undermines the positive risk-return relation with optimistic expectations, bad market news in low-sentiment periods strengthens the positive risk-return relation with pessimistic expectations. This result is consistent with the notion that high investor sentiment amplifies mispricing and is also naturally explained by uninformed investors' mis-reaction to price changes.[4] Fourth, the results for the business cycle indicate that, while the degree of optimistic (pessimistic) expectations is relatively stronger during expansion (recession) periods, they are not sufficiently different to have different impacts on the asymmetric intertemporal risk-return relation. The result implies that the attenuation (reinforcement) of the positive risk-return relation under good (bad) market news is robust to variation across phases of the business cycle. Finally, the availability of options expands the ability of

[4] In contrast to earlier studies that capture periods of mispricing through the low frequency investor sentiment data of Baker and Wurgler (2006), we find that mispricing caused by biased expectations in response to daily market news significantly affects the risk-return relation in the same manner as does investor sentiment.

informed investors to exploit mispricing and further reinforces the distortion of the positive risk-return relation. This result supports our conjecture that mispricing caused by the biased expectations of uninformed investors is exploited by rational arbitrageurs' selling on good news and buying on bad news, which ultimately distorts the positive intertemporal risk-return relation.

The outline of this paper is as follows. Section 2 reviews relevant literature. Section 3 presents our empirical work and interpretation of the results. Sections 4 and 5 present the empirical results on the effect of investor sentiment and the business cycles, respectively. Section 6 presents the empirical results on the effect of the availability of options. Section 7 presents our discussion concerning the asymmetric risk-return tradeoff. Section 8 concludes the paper.

2. **Literature Review**

A positive risk-return tradeoff is consistent with the time-varying rational expectations hypothesis in which time variation in the expected risk premium can be attributed to rational, risk-averse investors' revising their expectations in response to time-varying volatility. Supporting this conclusion, many studies have empirically documented a positive risk-return trade-off. For example, Pindyck (1984) shows that much of the decline in U.S. stock prices during the 1970s can be attributed to the upward shift in market risk premium arising from high market volatility and suggests that a substantial portion of time variation in the expected market risk premium is caused by time-varying risk factors in investment opportunities. Ghysel et al. (2005), and Bali (2008) present evidence that a substantial time variation in expected market returns is induced by a positive risk-return tradeoff. Using the implied cost of capital as a measure of expected market return, Pastor et al. (2008) document a positive risk-return relation in the G-7 stock markets. Employing both daily and intraday (5-minutes) return data, Bali and Peng (2006) find a positive risk-return relation. Guo and Neely (2008) use 30 years of daily data for 19 major international stock markets and find a significantly positive risk-return relation. Lundblad (2007) shows that a large data span instead of data frequency is required to reliably detect the risk-return relation and finds a significantly positive intertemporal risk-return relation for the sample period from 1836 to 2003 for both the U.S. and U.K. stock markets. Using a large panel, Wang et al. (2021) find a positive risk-return relation for 28 international stock markets. Other studies that support a positive intertemporal risk return relation include Haugen et al. (1991), Scruggs (1998), and Ludvigson and Ng (2006), to name a few.

On the other hand, while a positive intertemporal risk-return relation is consistent with theoretical predictions, some argue that the risk-return relation can be close to zero or even negative. For example, using an instrumental variable method with market interest rates, Campbell (1987) finds a significantly negative risk-return relation. Abel (1988) argues that a negative relation between conditional risk and the risk premium is consistent with a general equilibrium model when the coefficient of relative risk aversion is less than one. Barky (1989) also documents that the directional effect of an increase in risk on stock prices depends on the curvature of the utility function, which suggests the possibility of a negative risk-return relation. Glosten et al. (1993) argue that the sign of intertemporal risk return relation can be negative when investors are exceptionally optimistic about future stock price performance, thus not requiring a large premium for bearing risk. Brandt and Kang (2004) find that the conditional mean and volatility are negatively correlated contemporaneously but positively correlated unconditionally due to the positive lead-lag relation between the two moments of stock returns. Among others, Gennotte and Marsh (1993), Backus and Gregory (1993), Nelson (1991), Whitelaw (1994), and Ang et al. (2006) support a negative intertemporal relation. Additional empirical works fail to support a positive risk-return relation by documenting that it is either unstable or essentially zero, causing insignificant time variation in the expected market risk premium, including Poterba and Summers (1986), Baillie and DeGennaro (1990), Boudoukh et al. (1997), Whitelaw (2000), and Müller et al. (2011).

The aforementioned studies are mainly concerned with the estimation of the risk-return relation assuming a constant relative risk aversion (RRA) parameter. A growing body of literature has documented that the intertemporal risk-return relation is not time invariant, but rather differs across discrete states. Estimating conditional mean return and variance using the instrumental variable method, Whitelaw (1994) finds that the risk-return relation could be positive or negative due to the unstable correlation between conditional mean and variance over time. Yu and Yuan (2011) document that the sign of the risk-return relation is strongly positive during a high investor sentiment regime while it is close to zero during periods of low sentiment. They suggest that during the high-sentiment period, the increased presence and trading of sentiment investors distort a positive risk-return relation. Employing the MIDAS (MIxed DAta Sampling) model through a Markov-switching specification, Ghysels et al. (2014) show that the risk-return relation is characterized by time-varying regimes with opposite signs in different regimes. Suggesting that the lagged market return is an important control variable for reliable estimation of the risk-return relation, Marks and Nam (2018) show that the positive risk-return relation is weakened under good market news but is strengthened under bad market news.[5] Using institutional investor sentiment data, Wang (2018) shows that the positive risk-return relation is attenuated when institutional investor sentiment is bullish but is not distorted under bearish institutional investor sentiment, thereby concluding that institutional investors are also sentiment driven investors. These studies provide strong evidence that the RRA is not constant over time.

3. Empirical Results

3.1. The Data and Volatility Measure

We use the daily returns on the CRSP value- and equal-weighted index as the nominal market portfolio returns for the full period from January 1926-December 2019 (24,792 observations) and two sub-samples as a robustness check. The 1st sub-period spans the period Jan. 2, 1926 – Dec. 31, 1987, while the 2nd sub-period covers the period Apr. 2, 1951 – Dec. 31, 2019. The 1st sub-period is widely used in the financial economics literatures for the reason that the sample period includes both the Great Depression and the 1987 stock market crash. The 2nd sub-period reflects the post-Treasury Accord period,

[5] Kang, et al. (2019) and Kilic, et al. (2022) show that the asymmetric risk-return relation along with the price adjustment process induce the short-term momentum, which can be the source of technical trading profits. Chelikani, et al. (2022) also employ the same methods to examine the ICAPM on the industry portfolio returns.

which is characterized by the modern Federal Reserve system. In March 1951, the U.S. Treasury and the Federal Reserve reached an agreement to separate government debt management from monetary policy. The Treasury-Federal Reserve Accord paved the way for the effective control of monetary policy by the Federal Reserve as the nation's central bank. Note that using a part of this post-Treasury Accord period data (April 1951 – December 1989), Glosten, Jagannathan, and Runkle (1993) show that the period exhibits a negative intertemporal risk-return relation. To compute the daily excess returns, we employ the U.S. one-month Treasury bill rate reported by Ibbotson Associates as the risk-free rate. Since the risk-free rate is only available on a monthly frequency, we construct the daily risk-free rate by assuming that it is constant within a month. The daily excess portfolio return is the difference between the nominal daily portfolio return and the daily risk-free rate.

Table 1 presents descriptive statistics for the daily excess returns of the value- and equal-weighted market portfolios. The daily excess returns exhibit the commonly observed properties of negative skewness, excess kurtosis, and significant return persistence at short horizons. Table 1 shows that partial return autocorrelations are significantly positive up to the 4th or 5th order at the 5% level, and the sum of the return autocorrelation coefficients up to the 5th order, $\psi(1)$, is positive for all sample periods. Accordingly, in all of the following empirical models, we specify an AR(5) process to represent the price adjustments resulting from investor behavior to correct prior mispricing.

As the proxy for conditional market volatility, we employ the conditional variance of the daily excess returns ($\hat{\sigma}^2_{m,t+1}$) from estimating the EGARCH (1,2) model proposed by Nelson (1991).[6] For the estimation of the indirect risk-return relation, we employ volatility forecast errors computed as $e^2_{m,t+1} - \hat{\sigma}^2_{m,t+1}$. For the estimation of the intertemporal risk-return relation, we adopt a two-step method. The first

[6] The EGARCH (1,2) model is specified as $\ln(\hat{\sigma}^2_t) = w + \theta v_{t-1} + \gamma\left(|v_{t-1}| - \sqrt{2/\pi}\right) + g_1 \ln(\hat{\sigma}^2_t) + g_2 \ln(\hat{\sigma}^2_{t-1})$, where $v_{t-1} = e_{t-1}/\sqrt{\hat{\sigma}^2_{t-1}}$, and $\sqrt{2/\pi} = E|v_{t-1}|$, and e_{t-1} is the mean-deviated daily excess market return at time $t-1$. The selection of the EGARCH (1,2) model is based on the log likelihood-ratio (LR) test. The computed value of the LR test statistic between the EGARCH (1,1) model and the EGARCH (1,2) model is 27.60, which is greater than the table value (6.635) under the $\chi^2_{(df=1)}$ at the 1% significance level. The LR test indicates that the EGARCH (1,2) model is statistically better than the EGARCH (1,1) model.

step is to obtain the conditional variance ($\hat{\sigma}^2_{m,t+1}$) from the EGARCH (1,2) model. In the second step, using the estimated conditional variance as the proxy for conditional market volatility, we estimate the intertemporal risk-return relation by running ordinary least squares (OLS) regressions on the excess market returns. In estimating the intertemporal risk-return relation, this two-step approach outperforms the maximum likelihood estimation (MLE) of GARCH-in-Mean models. While the estimate of the RRA parameter from the two-step method is the best linear unbiased estimator, the MLE estimates of the GARCH-in-Mean are known to suffer from sensitivity to the choice of the conditional distribution.[7] We also confirm that the MLE estimates of the GARCH-in-Mean model are inconsistent and unreliable under different distributional assumptions in the context of the empirical models considered in this paper.

[Insert Table 1 about here]

3.2. The State-Dependent Risk-Return Relation under Optimistic and Pessimistic Expectations

Yu and Yuan (2011) and Marks and Nam (2018) document that a normally positive risk-return relation is distorted (reinforced) conditional on high (low) market sentiment and good (bad) market news. We conjecture that this state-dependent risk-return relation is caused by uninformed investors' tendency to mis-react to market conditions with optimistic and pessimistic expectations, which are exploited by the arbitrage of overpriced (underpriced) stocks by informed traders. To test our conjecture, we employ the models suggested by Marks and Nam (2018), which suggest that investors' behavior to correct prior mispricing is also an important contributor to the expected market risk premium. Following Marks and Nam (2018), we augment the intertemporal risk-return relation with an AR(5) process as a way of incorporating the price adjustment process induced by investors' behavior in the following three models.

[7] See, Baillie and DeGennaro (1990) for details.

Model 1:

$$r_{m,t+1} = c + \delta\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}, \tag{1}$$

Model 2:

$$r_{m,t+1} = c + (\delta_P Pd + \delta_N Nd)\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}, \tag{2}$$

Model 3:

$$r_{m,t+1} = (c_P Pd + c_N Nd) + (\delta_P Pd + \delta_N Nd)\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}, \tag{3}$$

where $r_{m,t+1}$ is the excess market returns as a proxy for the expected market risk premium and $\hat{\sigma}^2_{m,t+1}$ is the conditional forecast of market volatility. $Pd(Nd)$ is the dummy that captures prior d-day consecutive positive (negative) market returns. $Pd(Nd)$ is the dummy that captures prior d-day consecutive positive and negative returns. For $d = 2$, for example, $P2 = 1$ when $e_{m,t-1} > 0$ and $e_{m,t} > 0$ while $N2 = 1$ when $e_{m,t-1} < 0$ and $e_{m,t} < 0$, where $e_{m,t}$ is the mean-deviated excess market return. While δ is the *constant* RRA parameter, $\delta_P(\delta_N)$ is the RRA parameter under prior positive (negative) changes in the excess market returns. $\sum_{j=1}^{5}\phi_j r_{m,t+1-j}$ is an AR(5) process that captures the price adjustments. $\sum_{j=1}^{5}\phi_j > 0$ is defined as a price adjustment causing return persistence, while $\sum_{j=1}^{5}\phi_j < 0$ is defined as the price adjustment inducing return reversal.

We focus on three relations when estimating the above models. Note that these points are not discussed in Yu and Yuan (2011) and Marks and Nam (2018). First, we check if the *constant* RRA parameter (δ) is a biased estimate of the risk-return relation. The result of $\delta > \delta_P(\delta < \delta_N)$ implies that the *constant* RRA parameter overestimates (underestimates) the risk-return relation conditional on good (bad) market news. Second, we check if market imperfections measured by c_P and c_N are consistent with optimistic and pessimistic expectations under good and bad market news, respectively. We define the result of $c_P > 0(c_N < 0)$ as evidence of optimistic (pessimistic) expectations. The rationale for this definition is that since prior mispricing is partly corrected by the AR(5) process, c_P and c_N should capture the component caused by investors' potentially optimistic or pessimistic expectations under good and bad market news that remains unexplained by a state-dependent RRA parameter. Third, we interpret estimates of $c_P > 0$ and $c_N < 0$ together with $\delta_P \leq 0$ and $\delta_N > 0$ as evidence supporting our argument that arbitrageurs' short-selling (buying) of overpriced (underpriced) stocks attenuates (strengthens) the positive intertemporal risk-return relation under good (bad) market news.

The estimation results of Models 1-3 are reported in Table 2. There are several notable findings that strongly support our conjecture. First, the RRA parameter is significantly positive under a prior negative price change but is non-positive under a prior positive price change.[8] For example, when using 1-day prior positive and negative dummies in Model 2, δ_N (t-value) has a significantly positive estimate of 3.647 (2.39) and δ_P (t-value) is insignificantly negative at –1.392 (–0.87). A similar pattern is observed for the estimates when using one-standard deviation dummies in Model 2: δ_N (t-value) = 6.415 (4.47) and δ_P (t-value) = –2.737 (–1.48). The estimates for the equal-weighted portfolio reported in Appendix Table 1 also show the same results. For the 1-day prior positive and negative return dummies in Model 2, δ_N (t-value) is 5.725 (3.80) and δ_P (t-value) is 1.112 (0.87), while δ_N (t-value) is 7.613 (4.27) and δ_P (t-value) is –0.397 (–0.26) for the one-standard deviation positive and negative return dummies. Model 3 also yields similar results for both portfolios, indicating that the asymmetric risk-return relation under positive versus negative price changes is robust to the presence of asymmetric market imperfections. For the value-weighted portfolio, δ_N (t-value) is 4.029 (2.43) and δ_P (t-value) is –1.997 (–1.19) for the 1-day prior positive and negative return while δ_N (t-value) is 6.521 (4.40) and δ_P (t-value) is –2.547 (–1.26) for the one-standard deviation dummies. The estimation results imply that the positive risk-return relation is attenuated under good market news while it is strengthened under bad market news.

Second, the results for both portfolios show that the asymmetry in the risk-return relation is more pronounced for the dummies indicating 2 and 3 consecutive past daily price changes in the same direction relative to just a prior one-day price change. This suggests that mispricing caused by short-term mis-reaction is more severe after a sequence of similar returns, and hence the asymmetry resulting in the risk-

[8] We use 'non-positive' to signify 'not significantly positive'.

return relation arising from mispricing is greater under 2 and 3 consecutive positive and negative price changes. Third, the constant RRA parameter is a biased estimate of the intertemporal risk-return relation. The results of Models 2 and 3 for both portfolios indicate that, compared to the significantly positive value of δ in Model 1, the value of δ_P becomes negative while the value of δ_N is significantly positive. This implies that assuming a constant RRA parameter *overestimates* the risk-return relation conditional on good market news, while *underestimating* the relation conditional on bad market news. Fourth, the estimates of $\phi(1)$ are significantly positive at the 1% level in all cases, implying that the adjustment process resulting from investor behavior to correct prior mispricing leads to return persistence.[9]

Finally, the results from Model 3 show that for both market portfolios, $c_P(c_N)$ is significantly positive (negative) at the 1% level for all estimates except for the case with dummies indicating returns one standard deviation above and below the mean. As stated earlier, the significantly positive (negative) value of $c_P(c_N)$ is strong evidence of investors' optimistic (pessimistic) expectations about the future performance of stocks under good (bad) market news. Thus, our results of $c_P > 0$ and $c_N < 0$ support the notion that the tendency for uninformed investors to be optimistic (pessimistic) in response to good (bad) market news causes overpricing (underpricing). More importantly, the estimates of $\delta_P < 0$ and $c_P > 0$ ($\delta_N > 0$ and $c_N < 0$) support our argument that the attenuation (reinforcement) of the positive risk-return relation is associated with arbitrageurs' short-selling of overpriced stocks (buying underpriced stocks) resulting from uninformed investors' mispricing under optimistic (pessimistic) expectation.

[Insert Table 2 about here]

[9] We perform a robustness check to examine whether there is asymmetry in the autoregressive process that affects the observed distortion of a positive intertemporal risk-return relation. To do so, we attach the price dummies not only to the RRA parameter but also to the return autoregressive process. The estimation results indicate that there is *no* significant asymmetry in the AR(5) process under positive and negative price changes, and that allowing for such an asymmetry in the AR(5) process does not affect our main results of return persistence.

As a robustness check, we examine the effect of extreme positive and negative price changes on the RRA parameter by using the dummies representing 4-, 5-, 6-day consecutive positive and negative returns, and two standard deviation positive and negative return changes in Models 2 and 3. The results are presented in Table 3, which are very similar to those reported in Table 2; the RRA parameter is negative (positive) conditional on recent extreme positive (negative) returns. The results for the equal-weighted portfolio reported in Appendix Table 2 also show similar results.

[Insert Table 3 about here]

We also perform sub-period analysis to examine the robustness of our main results and the stability of our models' description of an asymmetric risk-return relation and the significance of investors' adjustment behavior leading to return persistence. Using two sub-samples, Jan. 2, 1926 – Dec. 31, 1987 (Pre-1987 Crash period) and Apr. 2, 1951 – Dec. 31, 2019 (the post-Treasury Accord period), we estimate Models 1, 2, and 3 to check consistency across the two subsamples for the value-weighted portfolio. Table 4 reports the estimation results of the two sub-periods, which indicate that the main results obtained from the full sample period are indeed present in both subsamples. Notable findings are as follows. First, for both subperiods the estimation results of Model 3 show that the value of $c_P(c_N)$ is significantly positive (negative) at the 1% level for all estimates except for the one standard deviation return dummies. The result of $c_P > 0(c_N < 0)$ is consistent with the result for the full period as evidence of optimistic (pessimistic) expectations under good (bad) market news. Second, the estimates of Models 2 and 3 show that the asymmetric risk-return relation documented in the full sample period is still significant in both sub-periods. For both sub-periods, the estimated RRA parameter is significantly positive (negative) under a prior negative (positive) return. Third, also consistent with the result obtained from the full sample period, the magnitude of the RRA coefficients is much greater following 2 and 3 consecutive returns in the same direction than under a prior one-day return, implying that the distortion of the risk-return relation is greater following these short trends. Lastly, the estimates of $c_P > 0$ and $\delta_P < 0$ together with

$c_N < 0$ and $\delta_N > 0$ supports our argument that the asymmetric risk-return relation is caused by arbitrageurs' short-selling (buying) of overpriced (underpriced) stocks under good (bad) market news.[10]

[Insert Table 4 about here]

3.3. Indirect Test of Risk-Return Relation

French, Schwert, and Stambaugh (1987) state that "… *a positive relation between the predicted stock market volatility and the expected risk premium induces a negative relation between the unpredicted component of volatility and excess holding period returns*." [pg.15] They examine this negative relation between excess market returns and the contemporaneous unexpected volatility as an indirect test of the positive risk-return relation. We perform this indirect test conditional on good and bad market news. Since our observed risk-return relation is significantly *positive* under bad market news, we expect a strong *negative* relation under bad market news in this indirect test. To examine our conjecture, we estimate the following Model 4:

Model 4:

$$r_{m,t+1} = c_1 P + c_2 N + (\pi_P P + \pi_N N)\hat{\eta}^2_{m,t+1} + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{p} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}, \quad (4)$$

where $\hat{\sigma}^2_{m,t+1}$ is the conditional market volatility, and $\hat{\eta}^2_{m,t+1}$ is the contemporaneous volatility innovation representing unexpected volatility changes, which we compute as $\hat{\eta}^2_{m,t+1} = e^2_{m,t+1} - \hat{\sigma}^2_{m,t+1}$. The indirect risk-return relation is measured by $\pi_P (\pi_N)$ under a prior positive (negative) return. The observed positive

[10] An anonymous referee has suggested that we examine the robustness of the main result to different GARCH models. We have re-estimated Model 3 using both the symmetric GARCH (1, 2) and the asymmetric TARCH (1, 2) models as a robustness check. The estimation results are presented in Appendix Table A.7, which shows a significant asymmetric risk-return relation in which the positive risk-return relation is attenuated (strengthened) under good (bad) market news. The results verify that our finding of the asymmetric risk-return relation is robust to the presence of different GARCH models.

risk-return relation under a prior negative return is consistent with a significantly negative value of π_N under a prior negative return.

The estimates of Model 4 are reported in Table 5. Panel A provides results for the full sample period, while Panels B and C present the results for the two sub-periods. The results in all three Panels show strong evidence to support the asymmetric intertemporal risk-return relation under positive and negative returns. The estimated value of π_N is economically and statistically significantly negative at the 5% level and is robust over various parameter restrictions. The result of the *negative* indirect risk-return relation under a prior negative return confirms that the ex-ante risk-return relation is positive under a bad market news. The estimated value of π_P is significantly positive or weakly negative. This result is also consistent with the observed weak or negative risk-return relation under good market news. In sum, the results from the indirect test on the relationship between excess returns and the contemporaneous unexpected volatility provide robust evidence to support our conclusion concerning the asymmetric risk-return relation.

[Insert Table 5 about here]

4. The Effect of Investor Sentiment on the Asymmetric Intertemporal Risk-Return Relation

Baker and Wurgler (2006) construct an index measuring aggregate investor sentiment and then show that the future returns of stocks susceptible to mispricing depend on the current level of sentiment. Using Baker and Wurgler's (2006) composite sentiment index, Yu and Yuan (2011) argue that greater participation of sentiment traders in high sentiment periods attenuates the positive intertemporal risk-return relation at these times. This occurs because high sentiment causes these investors to overprice stocks, but these investors are reluctant to take short positions when sentiment is low, and therefore the distortion of the risk-return relation caused by sentiment traders is concentrated mainly during periods of high sentiment. Using an empirical model based on anomalies to identify stocks that are likely to be overpriced or underpriced, Stambaugh, Yu, and Yuan (2015) show that in the cross-section, a negative risk-return relation among overpriced stocks is significantly stronger during high-sentiment periods while a positive risk-return relation among underpriced stocks is strengthened during low-sentiment periods.

In this section, we examine whether the level of investor sentiment causes more profound mispricing. If sentiment traders are in general less rational and are more likely to mis-react to price changes, and these traders are also more actively engaged in overpricing through more buying during periods of high sentiment, then the distortion of the positive risk-return tradeoff caused by short selling overpriced stocks under good market news should be greater when sentiment is high. Likewise, if sentiment traders are relatively more pessimistic in response to bad market news and hence actively engaged in underpricing through more selling during periods of low sentiment, then the positive risk-return relation that rational arbitrageurs cause through purchasing underpriced stocks conditional on bad market news should be strengthened when sentiment is low. We thus conjecture that the attenuation (reinforcement) of a positive risk-return relation under good (bad) market news is stronger in high (low) sentiment periods than in low (high) sentiment periods. Following Shen, Yu, and Zhao (2017), we define high and low sentiment periods based on the sign of the orthogonalized monthly sentiment index and construct the daily sentiment index by assuming that it is constant within a month.[11] We first estimate the constant RRA parameter under high and low market sentiment as the benchmark model:

Model 5: Constant RRA under high/low market sentiment

$$r_{m,t+1} = \left[c^H + \delta^H \hat{\sigma}^2_{m,t+1} + \phi_H(1) r_{m,t}\right] \cdot H + \left[c^L + \delta^L \hat{\sigma}^2_{m,t+1} + \phi_L(1) r_{m,t}\right] \cdot L + \varepsilon_{m,t+1}, \qquad (5)$$

where $\phi_H(1)$ and $\phi_L(1)$ are the sum of AR(5) coefficients in the high- and low-sentiment regimes, and H (L) is the dummy representing low- (high-) sentiment regimes over July 1965 – December 2018. We

[11] We employ the monthly composite sentiment index from July 1965 to December 2018. Using the monthly composite sentiment index, we construct the daily sentiment index by assuming that it is constant within a month. Also, we thank Jeffrey Wurgler for making the index data available at http://people.stern.nyu.edu/jwurgler/.

employ this sample period due to data availability of the index. The RRA parameter is measured by $\delta^H(\delta^L)$ in the high- (low-) sentiment regime. Second, we estimate the following Model 6 to examine the impact of market sentiment on the asymmetric risk-return relation under good and bad market news.

Model 6: Asymmetric Nonlinear RRA under high/low market sentiment

$$r_{m,t+1} = [c^H + (\delta_P^H Pd + \delta_N^H Nd)\hat{\sigma}_{m,t+1}^2 + \phi_H(1) r_{m,t}] \cdot H + [c^L + (\delta_P^L Pd + \delta_N^L Nd)\hat{\sigma}_{m,t+1}^2 + \phi_L(1) r_{m,t}] \cdot L + \varepsilon_{m,t+1}, \qquad (6)$$

where the RRA parameter in the high-sentiment regime is measured by $\delta_P^H(\delta_N^H)$ under prior *d*-day positive (negative) returns, while $\delta_P^L(\delta_N^L)$ measures the RRA parameter under prior *d*-day positive (negative) returns in the low-sentiment regime. We extend the dummies to the case of 3-day consecutive returns to examine whether a tendency to perceive trends in short samples of returns would induce even greater mispricing, and hence exhibit a greater distortion of the positive risk-return relation. We also use the dummies capturing prior one and two standard deviation returns.

The estimation results are reported in Table 6, which shows several notable findings. First, the results for Model 5 (the constant RRA) shows $\delta^H = 1.929$ (or 2.946) and $\delta^L = 1.520$ (or 0.980), with only δ^H significant at the 5% level. The results are a reference point for Model 6 (the asymmetric RRA) in examining if this positive risk-return relation is strengthened (weakened) under prior negative (positive) returns. Second, the estimated values of δ_P^H and δ_P^L are all negative except for the cases with 1-day return associated with insignificant positive values, while the estimated value of δ_N^H and δ_N^L are all positive. The average estimated value of the RRA coefficients across the 10 estimations is $\delta_P^H = -5.877$, $\delta_P^L = -4.401$, $\delta_N^H = 5.924$ and $\delta_N^L = 9.324$, implying that the asymmetric risk-return relation under good and bad market news is still robust to the presence of the investor sentiment.

Third, mispricing under 2- and 3-day consecutive price changes causes a greater impact on the risk-return relation when compared to 1-day return. For example, the estimated value of $\delta_P^H = 1.666$

under a 1-day return dramatically decreases to –6.562 and –11.486 for 2- and 3-day consecutive returns. Likewise, the estimated value of $\delta_N^L = 5.407$ under a 1-day price change increases to 10.040 and 14.379 for 2- and 3-day consecutive returns. While having three-consecutive same signs for daily returns is a very short trend, the results indicate that such sequences may be required to have significantly different effects on the uninformed investors' expectations. In particular, the tendency for sentiment-driven investors to mis-react to price changes would be amplified by consecutive returns of the same sign, and thus the differential effect of high versus low sentiment on the ex-ante risk-return relation could be more clearly observed by extending the length of the sequence. Thus, our results demonstrate that a greater number of consecutive returns amplifies mispricing and the distortion it causes on the RRA parameter.

Fourth, there is a significant differential effect of high versus low sentiment on the asymmetric RRA parameters. For all estimations, the value of δ_N^L is significantly positive at the 5% level and greater than δ_P^L. Also, except for the case with a prior one-day positive and negative return, the estimated value of δ_P^H is significantly negative at the 5% level and less than δ_P^L. If high sentiment causes greater mispricing and hence more short-selling, overpricing in response to good market news should attenuate the positive risk-return relation more in high-sentiment periods than in low-sentiment periods, while bad market news should strengthen the positive risk-return relation more in low-sentiment periods than in high-sentiment periods. Thus, the results of $0 > \delta_P^H < \delta_P^L$ and $0 < \delta_N^L > \delta_N^H$ support our conjecture that the attenuation (reinforcement) of a positive risk-return relation under good (bad) market news is stronger in high (low) sentiment periods than in low (high) sentiment periods.

To detect evidence of optimistic (pessimistic) expectations during high (low) sentiment periods, we estimate a variant of Model 6 with new constant terms specified as $(c_P^H + c_P^L)Pd + (c_N^H + c_N^L)Nd$ for $d = 1, 2, 3$. The estimates show a consistent result of $c_P^H > 0$, $c_P^L > 0$, $c_N^H < 0$, and $c_N^L < 0$ with $c_P^H > c_P^L$ and $|c_N^H| < |c_N^L|$. While the results of $c_P^L > 0$ and $c_N^L < 0$ are significant at the 1% level for all three estimations, the result of $c_P^H > 0$ is significant at the 5% level for $d = 2, 3$, and the result of $c_N^H < 0$ is significant at the 5% level $d = 1, 3$. The results of $c_P^H > c_P^L > 0$ and $c_N^L < c_N^H < 0$ indicate that, while good (bad) market news consistently causes optimistic (pessimistic) expectations, the impact of optimistic (pessimistic) expectations is relatively stronger during high (low) sentiment periods.[12] The results for the equal-weighted portfolio reported in Appendix Table 4 also show similar results.

[Insert Table 6 about here]

While Baker and Wurgler's index is a commonly used proxy in sentiment literature, several studies employ alternative proxies to test the effect of sentiment on expected return. For example, Schmeling (2009) and Lemmon and Portniaguina (2006) use the consumer confidence index to analyze the effects of investor sentiment on the international and US stock markets. Constructing the FEARS (Financial and Economic Attitudes Revealed by Search) index as a new measure of investor sentiment, Da, Engelberg, and Gao (2015) show that their results are consistent with theories of investor sentiment. Creating a new investor sentiment index by eliminating a common noise component in sentiment proxies, Huang et al., (2015) show that the PLS (Partial Least Squares) index has outperform the existing sentiment indices in terms of predictive power both in and out of sample. Constructing the AS (Augmented Sentiment) index from extracting the common information from the Baker and Wurgler index, the University of Michigan Consumer Sentiment Index, and Conference Board Consumer Confidence Index, Doukas and Han (2021) examine the effect of investor sentiment on the slope of the security market line. They show that the security market line implied by the sentiment-scaled beta exhibits an upward (downward) slope during high-sentiment (low-sentiment) periods. Examining the impact of both retail and institutional investor sentiments on the distortion of the positive risk-return relation, Duxbury and Wang (2023) show that their findings are robust to alternative sentiment proxies.

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[12] The estimated values of $c_P^H \times 100$ (t-value) for $d = 1, 2, 3$ are 0.026 (1.27), 0.125 (3.17), and 0.079 (2.20), while the values of $c_P^L \times 100$ (t-value) are 0.114 (5.24), 0.090 (2.53), and 0.100 (4.17). The estimated values of $c_N^H \times 100$ (t-value) are -0.055 (-2.14), -0.042 (-1.20), and -0.061 (-1.97), while the values of $c_N^L \times 100$ (t-value) are -0.066 (-3.11), -0.139 (-3.56), and -0.112 (-2.21). The estimated values of RRA parameters are almost same as those shown in Table 6 with some variations. We omit the full results for brevity.

Employing the same proxies that are used in Doukas and Han (2021), we conduct the robustness check of our result to the alternative measures of investor sentiment.[13] The results of the robustness check are presented in Table 7. While Panel A shows the results of the robustness check for the University of Michigan Consumer Sentiment Index, Panel B presents the results for the Augmented Sentiment index as the proxy for sentiment. Robustness.[14] The results show strong evidence to support the asymmetrical effect of sentiment on the expected market risk premium, which is consistent with the result of Doukas and Han (2021). The results in both panels show that, while the estimates of δ_N^H and δ_N^L are all positive, those of δ_P^H and δ_P^L are either negative or negligible, verifying that the attenuation (reinforcement) of a positive risk-return relation under good (bad) market news is stronger in high (low) sentiment periods than in low (high) sentiment periods. In sum, the results of the robustness check imply that the asymmetric risk-return tradeoff under good and bad market news is still robust to the presence of the alternative proxies for the high- and low-sentiment regimes.

[Insert Table 7 about here]

5. The Impact of Business Cycle on the Asymmetric Intertemporal Risk-Return Relation

The channels which link stock market movements to the business cycle have been studied extensively. For example, Chen, et al. (1986), Keim and Stambaugh (1986), Fama (1990), Schwert (1990), and Chen (1991) document that fluctuations in the level of economic activity are a key determinant of the level of stock returns. Fama and French (1989) show that expected returns include risk

[13] An anonymous referee has suggested that we perform the robustness check of our result using alternative proxies for investor sentiment. We appreciate the referee's suggestion. Following Doukas and Han (2021), we employ the University of Michigan Consumer Sentiment Index, the Consumer Confidence Index, and the Augmented Sentiment index for the robustness check.

[14] For space reasons, the estimation results for the Consumer Confidence Index are not included in this paper but are available upon request.

premiums that move inversely with business conditions. Whitelaw (1994) reports that expected returns move inversely with the business cycle and conditional volatility leads the expected return cycle. Gibson and Mougeot (2004), Chordia et al. (2005), and Næs et al. (2011) find a strong relation between stock market liquidity and the business cycle. Campbell and Diebold (2005) find that expected business conditions forecast higher market risk and hence affect the expected market risk premium. Choudhry et al. (2016) find that there is a bidirectional causal relationship between stock market volatility and the business cycle. Ghysels et al. (2005), Bali (2008), Kim and Lee (2008), Bali and Engle (2010), Nyberg (2012), among others, study how the business cycle affects the risk-return tradeoff.[15] These studies use macroeconomic state variables as the proxies for investment opportunities defined in Merton (1973).

Given the fact that there is a significant relationship between time-varying stock market volatility and fluctuations in the level of real economic activity, we examine whether and how the business cycle affects uninformed investors' optimistic and pessimistic expectations and the resulting distortion of the positive risk-return relation under good and bad market news. Using the National Bureau of Economic Research (NBER) business cycle indicator, we generate dummies to represent the state of economy in terms of expansion and recession periods. We first estimate the following constant RRA model under expansion and recession periods as the benchmark model:

Model 7: Constant RRA under expansion/recession periods

$$r_{m,t+1} = \left[\mu^E + \delta^E \hat{\sigma}^2_{m,t+1}\right] \cdot E + \left[\mu^R + \delta^R \hat{\sigma}^2_{m,t+1}\right] \cdot R + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}, \quad (7)$$

where $E(R)$ is a dummy representing an expansionary (recessionary) period defined by the NBER. Periods of expansion begin at the trough date and end at the peak date, while periods of recession begin at

[15] While Kim and Lee (2008) find evidence on the positive relation only in the expansion regime, Nyberg (2012) finds a positive intertemporal risk-return relation independent of the business cycle.

the peak date and end at the trough date. We then estimate the following model to examine the effect of the business cycle on the asymmetric risk-return relation conditional on good and bad market news:

Model 8: Asymmetric Nonlinear RRA under expansion/recession periods

$$r_{m,t+1}\left[(\mu_P^E Pd + \mu_N^E Nd) + (\delta_P^E Pd + \delta_N^E Nd)\hat{\sigma}_{m,t+1}^2\right] \cdot E + \left[(\mu_P^R Pd + \mu_N^R Nd) + (\delta_P^R Pd + \delta_N^R Nd)\hat{\sigma}_{m,t+1}^2\right] \cdot R + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}, \qquad (8)$$

where $t$he RRA parameter in the expansion periods is measured by $\delta_P^E\,(\delta_N^E)$ under prior positive (negative) price changes, while $\delta_P^R\,(\delta_N^R)$ measures the RRA parameter under prior positive (negative) price change in the recession periods.

The estimation results for the value-weighted portfolio are reported in Table 8. There are two notable findings. First, for $d = 1, 2, 3$ the estimates show a consistent result of $\mu_P^E > 0$, $\mu_P^R > 0$, $\mu_N^E < 0$, and $\mu_N^R < 0$ with $\mu_P^E > \mu_P^R$ and $\left|\mu_N^E\right| < |\mu_N^R|$. While the results of $\mu_P^E > 0$ and $\mu_N^R < 0$ are significant at the 1% level for all three estimations, the result of $\mu_P^R > 0$ is significant at the 10% level for $d = 1, 2$, and the result of $\mu_N^E < 0$ is significant at the 5% level $d = 1, 3$. The average value of the estimates for the 5 estimations are $\mu_P^E = 0.056$, $\mu_P^R = 0.011$, $\mu_N^E = -0.018$, and $\mu_N^R = -0.193$. The results of $\mu_P^E > \mu_P^R > 0$ and $\mu_N^R < \mu_N^E < 0$ indicate that, while good (bad) market news consistently induces optimistic (pessimistic) expectations, the impact of optimistic (pessimistic) expectations seems relatively stronger during expansion periods.

Second, the estimated values of both δ_P^E and δ_P^R are either weakly or significantly negative for all estimations except for the value of δ_N^E with $d = 1$. The average estimated values of the RRA coefficient under prior positive price changes are $\delta_P^E = -4.039$ and $\delta_P^R = -4.414$, respectively, implying that the distortion of the positive risk-return relation under good market news is independent of the business cycle. Also, the estimated values of δ_N^E and δ_N^R are all positive and statistically significant at the 5% level for all estimations. The average estimated values of the RRA coefficients under prior negative price changes are

$\delta_N^E = 9.016$ and $\delta_N^R = 7.713$, and their difference is not statistically significant. This implies that the reinforcement of the positive risk-return relation under bad market news is still significant and appears independent of the business cycle. The estimation results for the equal-weighted portfolio presented in Appendix Table (A.5) also show similar results; the average values of the estimates are $\mu_P^E = 0.059$, $\mu_P^R = 0.096$, $\mu_N^E = 0.028$, and $\mu_N^R = -0.199$, while the average estimated value of the RRA coefficients are $\delta_P^E = -1.810$, $\delta_P^R = 0.829$, $\delta_N^E = 9.978$ and $\delta_N^R = 9.114$.[16]

In sum, the estimation results concerning the asymmetric risk-return relation conditional on the business cycle imply that the degree of optimistic (pessimistic) expectations seems relatively stronger during expansion (recession) periods, but the difference is not sufficient to induce a differential impact on the asymmetry in the RRA parameters between the expansionary and recessionary cycles. Rather, the results imply that the attenuation (reinforcement) of the positive risk-return relation under good (bad) market news is robust to the presence of the business cycle.

[Insert Table 8 about here]

6. The Short-Sale Effect of the Introduction of Options on Intertemporal Risk-Return Relation

6.1. Introduction of Options

Options offer strategic alternatives to informed investors. Many studies have documented that stock options can facilitate short positions, thereby improving the efficiency of financial markets. While both options traders and short sellers may be informed investors attempting to capture arbitrage profits, options typically offer an easier and less expensive way of taking short positions, allowing investors better access to short positions. In this case, options trades such as writing calls or holding long puts can

[16] There are some studies that control the correlation between investor sentiments and business cycles. Shen et al. (2017) show that over half of recession months are categorized as high-sentiment months, while Wang and Duxbury (2021) find that the correlation between the economic conditions and institutional investor sentiment is only at a medium level. However, our daily data of market sentiments and business cycles exhibits almost zero correlation between them.

function as substitutes to short selling equities. Sorescu (2000) and Danielsen and Sorescu (2000) report that the introduction of stock options caused the prices of the underlying stocks to fall in the period during 1981 – 1995, thus supporting the idea that options allow negative information to be better reflected in the stock prices. Showing a positive association between the level of short interest and stock lending fees, Boehme et al. (2006) suggest that bearish option trading and short selling are substitutes for each other. Huang et al. (2019) show that there exists a substitution effect between short selling and options trading in predicting aggregate stock returns. However, Battalio and Schultz (2006), Blau and Brough (2015), and DeLisle, et al. (2016) suggest that a causal link between more active options trading and more short selling indicates that they are complements to each other. Figlewski and Webb (1993), however, suggest that options trading plays a substitution role in mitigating the negative impact of short sale constraints, and options trading and short sales are also complementary to each other for hedging purposes.

The aforementioned studies imply that short selling and options trading are commonly used by informed traders to earn arbitrage profits so that the demand for options to implement short positions is substantial. This implies that there is a positive relation between option trading and stock trading. We thus conjecture that, due to the complementary and/or substitution effect of options trading on short selling or buying stocks, the distortion of the positive risk-return relation under good and bad market news should be stronger for the period after the introduction of options than prior to their availability. To examine our conjecture, we estimate the following models with the dummy variable (O) specified for the introduction of options into the equity markets:

Model 9A:

$$r_{m,t+1} = c + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta^O_P P + \delta^O_N N)\hat{\sigma}^2_{m,t+1} \cdot O + \sum_{j=1}^{5} \phi_j r_{m,t-j} + \varepsilon_{m,t+1}, \tag{9}$$

Model 9B:

$$r_{m,t+1} = (c_1 + c_2 O) + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta^O_P P + \delta^O_N N)\hat{\sigma}^2_{m,t+1} \cdot O + \sum_{j=1}^{5} \phi_j r_{m,t-j} + \varepsilon_{m,t+1}, \tag{10}$$

Model 9C:

$$r_{m,t+1} = (c_1 P + c_2 N) + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta^O_P P + \delta^O_N N)\hat{\sigma}^2_{m,t+1} \cdot O + \sum_{j=1}^{5} \phi_j r_{m,t-j} + \varepsilon_{m,t+1}, \quad (11)$$

where O is the dummy variable to represent the period after the introduction of options, i.e., January 1981 – December 2019. To define the dummies of $P(N)$, as before we consider three different cases of state-dependent price dynamics. For the first case, $P(N)$ represents a prior 1-day positive (negative) return. For the second and third cases, the dummies capture prior one or two standard deviation positive (negative) return changes, respectively. The RRA parameter after the introduction of options is measured by $\delta_P + \delta^O_P (\delta_N + \delta^O_N)$ under prior positive (negative) returns, such that $\delta^O_P (\delta^O_N)$ measures the differential effect of the availability of options on the risk-return relation under prior positive (negative) returns. We expect that the estimated value of δ^O_P and $\delta_P + \delta^O_P$ (δ^O_N and $\delta_N + \delta^O_N$) is significantly negative (positive).

The estimates of the three models for the value-weighted market portfolio are reported in Table 9, which support our conjecture concerning the short-sale effect on the distortion of the positive risk-return relation. First, the estimated value of δ^O_P is negative and statistically significant at the 5% level in all cases, verifying the significant short-sale effect of options trading conditioned on good market news. For all estimations, the estimated value of $\delta_P + \delta^O_P$ is also negative and statistically significant at the 5% level. The average value of the RRA coefficients with t-value across the 9 estimations is $\delta^O_P = -5.556\ (-2.15)$ and $\delta_P + \delta^O_P = -5.177\ (-2.79)$. The significant results of $\delta^O_P < 0$ and $\delta_P + \delta^O_P < 0$ imply that the distortion of the positive relation under good market news is more profound after the introduction of options than before. Second, the estimated value of δ^O_N is positive and statistically significant at the 5% level except for the two estimation results with the dummies of two standard deviation returns, which show a significance at the 10% level. The estimated value of $\delta_N + \delta^O_N$ is also positive and statistically significant at the 5% level, implying that options trading conditioned on bad market news has a significant positive effect on the risk-return tradeoff.[17] The average value of the RRA coefficients with t-value across the 9 estimations is $\delta_N^O = 4.704$ (2.01) and $\delta_N + \delta_N^O = 8.054$ (3.82). Third, we perform the two diagnostic tests, the F- and LR-tests, on the effect of options in the asymmetric risk-return relation. The results of the two diagnostic tests on the joint nulls of $\delta_P^O = 0$ and $\delta_N^O = 0$ verify that the introduction of options causes an economically and statistically significant effect on the asymmetric intertemporal risk-return relation. In sum, the availability of options appears to reinforce the ability of informed investors to exploit mispricing, with the statistically significant results of $\delta_P^O < 0$ and $\delta_N^O > 0$ with $\delta_P + \delta_P^O < 0$ and $\delta_N + \delta_N^O > 0$ supporting our argument that arbitrageurs' short selling of overpriced stocks under good market news (buying of underpriced stocks under bad market news) weakens (strengthens) the positive risk-return relation.[18]

[Insert Table 9 about here]

6.2. Placebo Test

The observed effect of the introduction of options on the risk-return tradeoff might be reflecting the continuation of pre-existing effect before the introduction of options. To ensure that the introduction of options induces the short-selling effect on the asymmetric risk-return tradeoff, we employ a placebo test.[19] Specifying the 10-, 20-, and 30-years periods before the introduction of options (i.e., January 1981)

[17] We also estimated the models for prior 2- and 3-days consecutive returns. The estimations yield the same sign of the RRA parameters as implied in our conjecture, with δ_P^O and δ_N^O insignificant. Interestingly, the estimated value of δ_P is significantly negative at the 5% level for all estimations. The average estimated value of δ_P (t-value) across the 6 estimations is –8.269 (–2.37). This result indicates that for the consecutive positive returns, the direct effect of short selling dominates the effect of options trading in the distortion of positive risk-return relation. The results are not presented here due to space limitations but are available upon request.

[18] The results for the equal-weighted portfolio presented in Appendix Table (A.6) also show similar results that document the short-sale effect on the distortion of the positive risk-return relation. The average value of the RRA coefficients with t-value across the 9 estimations is $\delta_P^O = -8.884$ (-4.36), $\delta_P + \delta_P^O = -6.864$ (-5.45), $\delta_N^O = 2.059$ (0.55), and $\delta_N + \delta_N^O = 8.698$ (2.18).

[19] The same anonymous referee has suggested that we perform a placebo test on the short-selling effect of the introduction of options on the intertemporal risk-return relation. We appreciate the referee's suggestion.

as the three randomized placebo periods, we examine whether the placebo effect causes the asymmetric intertemporal risk-return relation in the following three models:

Model 10A:

$$r_{m,t+1} = c + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta_P^{CB} P + \delta_N^{CB} N)\hat{\sigma}^2_{m,t+1} \cdot CB + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}, \quad (12)$$

Model 10B:

$$r_{m,t+1} = (c_1 + c_2 CB) + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta_P^{CB} P + \delta_N^{CB} N)\hat{\sigma}^2_{m,t+1} \cdot CB + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}, \quad (13)$$

Model 10C:

$$r_{m,t+1} = (c_1 P + c_2 N) + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta_P^{CB} P + \delta_N^{CB} N)\hat{\sigma}^2_{m,t+1} \cdot CB + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}, \quad (14)$$

where CB is the dummy variable to represent the three randomized placebo periods before the introduction of options into the markets. δ_P^{CB} and δ_N^{CB} are the RRA parameters to measure the placebo effect on the asymmetric intertemporal risk-return relation. If the observed short-selling effect of the introduction of options is caused by the placebo effect, then we will observe the economically and statistically significant value of $\delta_P^{CB} \leq 0$ and $\delta_N^{CB} > 0$.

The estimation results are reported in Table 10. Panels A and B respectively show the results for testing the effect of the 10- and 20-year placebo periods on the asymmetric risk-return relation. The estimation results in both panels show an economically and statically significant value of $\delta_P^{CB} > 0$ and $\delta_N^{CB} < 0$ and the estimated signs are completely opposite to those expected under the placebo effect. The average values of the placebo RRA coefficients across the 6 estimations are δ_P^{CB} (Newey-West *t*-value) = 13.923 (3.39) and $\delta_N^{CB} =- 14.084$ (-4.57) for Panel A, while they are $\delta_P^{CB} = 16.197$ (5.11) and $\delta_N^{CB} =$

-12.496 (-4.93) for Panel B. The testing results for the 30-year placebo period also show similar estimates. The average values of the placebo RRA coefficients across the 6 estimations are $\delta_P^{CB} = 15.315$ (5.28) and $\delta_N^{CB} =- 7.420$ (-2.14).[20] The results imply that there is no placebo effect on the short-selling effect of the introduction of options, and hence verifying that the introduction of options induces a significant short-selling effect on the asymmetric intertemporal risk-return relation.

[Insert Table 10 about here]

7. Discussion of Our Results

Suppose all stocks in the market are in equilibrium such that they are fairly priced in accordance with their relevant risk level and the conditional information concerning their fundamentals. In the simple case, good market news may take the form of a general upward revision in expected cash flows, a downward revision in the estimate of future risk, or some combination of the two that leads to a higher price level. Short-term uninformed investors who mis-react to good market news may become overly optimistic about future cash flows and/or may forecast future risk to be too low, resulting in a new price level that is too high. Once overpriced, the future return must be relatively low to restore the correct price level, so that as shown in Stambaugh, et al. (2015), rational arbitrageurs will try to profit from overpricing through short selling. As more arbitrageurs sell short overpriced stocks, the ex-ante risk-return relation observed at the time of overpricing will be lower than it would have been if there were no arbitrage. If the risk-return tradeoff is positive as standard asset pricing theory predicts, then at times of overpricing, short sales by arbitrageurs would tend to weaken the relation and make it less positive. Similarly, negative market news can take the form of a downward revision in expected cash flows and/or an increase in

[20] For space reasons, however, the testing results for the 30-year placebo period are not reported in this paper but are available upon request.

expected risk resulting in a lower price level. Mispricing resulting from negative market news implies underpricing due to pessimistic revisions in expected cash flows that are too negative and/or increases in expected risk that are too high. Rational arbitrageurs will exploit underpricing through purchasing underpriced stocks that produce relatively high future returns. Thus, at the time of underpricing the ex-ante risk-return tradeoff would tend to be more positive than it would otherwise be absent arbitrage.

Our results suggest that uninformed investors tend to mis-react to market news in the short-term, which can lead to the mispricing of stocks, such as overpricing conditional on good news but underpricing conditional on bad news. While price levels in the short-term may be distorted by investors with bounded rationality, arbitrageurs who are assumed to be fully rational could take advantage of mispricing. We argue that the observed asymmetric intertemporal risk-return relation is a consequence of rational arbitrageurs' trading to exploit mispricing through the selling of overpriced stocks conditional on good news and buying underpriced stocks conditional on bad news.

8. Conclusion

We further explore the state dependency in the risk-return relation by examining how the positive risk-return relation is distorted in response to various market conditions, such as extreme price changes, different levels of investor sentiment, the introduction of stock options, and phases of the business cycle. The patterns of state dependency that we document are indicative of the tendency for uninformed investors to mis-react to price changes in the short-term, thereby causing overpricing (underpricing) as a result of optimistic (pessimistic) expectations under good (bad) news. This mispricing is exploited by rational arbitrageurs' short selling of overpriced stocks and purchase of underpriced stocks, which distorts the risk-return tradeoff. We demonstrate that due to arbitrage trading, overpricing (underpricing) weakens (strengthens) the positive intertemporal risk-return relation, such that it is strongly positive conditional on bad market news, but non-positive conditional on good market news.

We study how the distortion of the relative risk aversion parameter varies across high and low sentiment periods. We find that while good market news in high-sentiment periods undermines the positive risk-return relation with optimistic expectations, bad market news in low-sentiment periods strengthens the positive risk-return relation with pessimistic expectations. This result is consistent with the notion that high investor sentiment amplifies mispricing. The pattern of a weak or negative risk-return relation following positive news or high market-sentiment, and a strongly positive risk-return relation following negative news or low market-sentiment, is naturally explained by uninformed investors' mis-reaction to price changes. Therefore, we conclude that investor mis-reaction to daily price changes due to optimistic and pessimistic expectations significantly impacts asset prices by causing mispricing, which ultimately attenuates or reinforces a typically positive ex-ante risk-return relation.

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Table 1
Descriptive Statistics

	Mean (×100)	STDV	SKEW	KURT	$\rho(1)$	$\rho(2)$	$\rho(3)$	$\rho(4)$	$\rho(5)$	$\psi(1)$
VWRETD	0.029	0.011	-0.123	19.734	0.069	-0.040	0.003	0.024	0.001	0.057
(Full-period)	(4.39)				(10.86)	(-6.30)	(0.47)	(3.78)	(0.16)	(8.97)
VWRETD	0.026	0.011	-0.074	23.724	0.113	-0.041	0.002	0.043	0.025	0.142
(1st Sub-period)	(3.24)				(14.62)	(-5.30)	(0.26)	(5.56)	(3.23)	(18.37)
VWRETD	0.029	0.009	-0.565	19.569	0.063	-0.030	0.008	-0.001	-0.015	0.025
(2nd Sub-period)	(4.13)				(8.30)	(-3.95)	(1.05)	(-0.13)	(-1.98)	(3.29)
EWRETD	0.071	0.010	0.336	28.785	0.213	-0.041	0.061	0.044	0.025	0.302
(Full-period)	(10.60)				(33.54)	(-6.46)	(9.60)	(6.93)	(3.94)	(47.55)
EWRETD	0.070	0.011	0.546	31.635	0.250	-0.074	0.074	0.048	0.038	0.336
(1st Sub-period)	(8.25)				(32.33)	(-9.57)	(9.57)	(6.21)	(4.91)	(43.46)
EWRETD	0.056	0.009	-0.621	16.539	0.201	0.017	0.061	0.046	0.017	0.342
(2nd Sub-period)	(9.07)				(26.47)	(2.24)	(8.03)	(6.06)	(2.24)	(45.04)

Note: This table reports descriptive statistics of daily excess returns of value- and equal-weighted market returns (VWRETD and EWRETD) of the NYSE, AMEX, and NASDAQ stocks retrieved from the CRSP data files for the full sample period (Jan. 2, 1926 – Dec. 31, 2019) and two sub-periods. The 1st sub-period spans the period Jan. 2, 1926 – Dec. 31, 1987, while the 2nd sub-period covers the period Apr. 2, 1951 – Dec. 31, 2019 (the post-Treasury Accord period). Daily excess market returns are computed by subtracting the daily average of monthly Treasury bill returns reported by Ibbotson Associates from the daily nominal returns of the market portfolios. STDV refers to the standard deviation. SKEW is skewness and KURT is kurtosis. $\rho(j)$ is the return autocorrelation coefficients at lag j. $\psi(1)$ is the sum of the five autocorrelation coefficients, i.e., $\psi(1) = \sum_{j=1}^{5} \rho(j)$. The numbers in parentheses below Mean are the t-values for the null of zero mean return, while those below $\rho(j)$ and $\psi(1)$ are the t-values computed with the Bartlett standard error $(1/\sqrt{N})$.

Table 2

Estimation Results of the Asymmetric Intertemporal Relation between Excess Market Returns and Conditional Market Volatility

	Constant RRA Parameter	Prior 1-day Positive/Negative Price Changes		Prior 2-day Consecutive Positive/Negative Price Changes		Prior 3-day Consecutive Positive/Negative Price Changes		One Standard Deviation Positive/Negative Return Changes	
	Model 1	Model 2	Model 3	Model 2	Model 3	Model 2	Model 3	Model 2	Model 3
$c_{(P)}(\times 100)$	0.012	0.012	0.080	0.016	0.091	0.016	0.084	0.015	0.015
	(1.61)	(1.70)	(5.86)	(1.50)	(5.71)	(1.81)	(4.30)	(2.14)	(0.29)
$c_N(\times 100)$			-0.058		-0.071		-0.111		-0.018
			(-3.84)		(-2.88)		(-3.64)		(-0.34)
$\delta_{(P)}$	1.468	-1.392	-1.997	-6.858	-7.587	-9.229	-11.068	-2.737	-2.547
	(2.19)	(-0.87)	(-1.19)	(-2.76)	(-2.93)	(-3.72)	(-3.95)	(-1.48)	(-1.26)
δ_N		3.647	4.029	7.760	8.280	11.137	13.180	6.145	6.521
		(2.39)	(2.43)	(3.92)	(3.81)	(4.03)	(4.18)	(4.47)	(4.40)
$\phi(1)$	0.067	0.099	0.062	0.155	0.112	0.155	0.128	0.117	0.116
	(2.86)	(4.38)	(2.51)	(5.57)	(3.73)	(4.88)	(3.94)	(5.05)	(4.15)
$AdjR^2(\%)$	0.75	0.83	1.08	1.14	1.30	1.30	1.42	1.02	1.00

Note: This table reports the estimates of the following models for the value-weighted market portfolio for the full period Jan. 2, 1926 – Dec. 31, 2019:

Model 1: $r_{m,t+1} = c + \delta\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 2: $r_{m,t+1} = c + (\delta_P Pd + \delta_N Nd)\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 3: $r_{m,t+1} = (c_P Pd + c_N Nd) + (\delta_P Pd + \delta_N Nd)\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

where $r_{m,t+1}$ is daily realized excess return of the value-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance estimated from the EGARCH (1, 2) model. ϕ_j is the j^{th} order return autocorrelation coefficient, and $\phi(1)$ is the sum of autocorrelation coefficients, i.e., $\phi(1) = \sum_{j=1}^{5}\phi_j$. $Pd(Nd)$ is the dummy to capture prior *d*-day positive (negative) returns, such that $P2 = 1$ when $e_{m,t-1} > 0$ and $e_{m,t} > 0$ ($N2 = 1$ when $e_{m,t-1} < 0$ and $e_{m,t} < 0$) where $e_{m,t}$ is the mean-deviated excess market returns. The RRA parameter is measured by $\delta_P(\delta_N)$ under prior *d*-day positive (negative) returns. We also estimate δ_P and δ_N using one standard deviation positive/negative return changes as the dummy variables. The numbers in parentheses are the Newey-West (1987) adjusted *t*-statistics. $Adj.R^2$ (%) is the percentage adjusted R^2.

Table 3
Estimation Results of the Asymmetric Intertemporal Risk-Return Relation Under Extreme Price Changes

	Prior 4-day Positive/Negative Price Changes		Prior 5-day Consecutive Positive/Negative Price Changes		Prior 6-day Consecutive Positive/Negative Price Changes		Two Standard Deviation Positive/Negative Return Changes	
	Model 2	Model 3	Model 2	Model 3	Model 2	Model 3	Model 2	Model 3
$c_{(P)}(\times 100)$	0.025	0.117	0.027	0.192	0.025	0.124	0.019	-0.056
	(3.39)	(4.86)	(3.57)	(2.98)	(3.63)	(4.12)	(2.87)	(-0.50)
$c_N(\times 100)$		-0.089		-0.061		0.037		0.061
		(-1.79)		(-0.88)		(0.44)		(0.55)
$\delta_{(P)}$	-15.214	-20.052	-21.869	-36.134	-6.528	-21.672	-1.512	-0.846
	(-3.66)	(-4.35)	(-1.86)	(-2.58)	(-1.11)	(-2.78)	(-0.71)	(-0.32)
δ_N	10.752	12.881	11.920	13.868	12.081	11.844	5.834	5.653
	(3.35)	(3.25)	(3.40)	(2.89)	(2.83)	(2.45)	(3.27)	(2.60)
$\phi(1)$	0.124	0.109	0.104	0.095	0.073	0.074	0.097	0.106
	(4.52)	(3.82)	(4.49)	(3.86)	(2.89)	(2.91)	(4.19)	(4.48)
$AdjR^2(\%)$	1.12	1.17	0.98	1.03	0.81	0.77	0.97	0.94

Note: This table reports the estimates of the following models for the value-weighted market portfolio for the full period Jan. 2, 1926 – Dec. 31, 2019:

Model 2: $r_{m,t+1} = c + (\delta_P Pd + \delta_N Nd)\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 3: $r_{m,t+1} = (c_P Pd + c_N Nd) + (\delta_P Pd + \delta_N Nd)\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

where $r_{m,t+1}$ is daily realized excess return of the value-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance estimated from the EGARCH (1, 2) model. $Pd(Nd)$ is the dummy to capture prior *extreme* *d*-day positive (negative) returns, i.e., *d* = 4, 5, 6. We also estimate the models using two standard deviation positive/negative return changes as the dummy variables. The RRA parameter is measured by $\delta_P(\delta_N)$ under prior *extreme* positive (negative) returns.

Table 4

Estimation Results of Sub-periods Analysis for the Asymmetric Intertemporal Risk-Return Relation

	Constant RRA	Prior 1-day Positive/Negative Price Changes		Prior 2-day Consecutive Positive/Negative Price Changes		Prior 3-day Consecutive Positive/Negative Price Changes		One Standard Deviation Positive/Negative Return Changes	
	Model 1	Model 2	Model 3	Model 2	Model 3	Model 2	Model 3	Model 2	Model 3
Panel A. 1st Sub-period: Jan. 2, 1926 – Dec. 31, 1987									
$c_{(P)}(\times 100)$	0.008	0.009	0.082	0.020	0.094	0.009	0.088	0.012	0.007
	(1.00)	(1.13)	(6.27)	(2.18)	(4.74)	(0.82)	(3.73)	(1.55)	(0.12)
$c_N(\times 100)$			-0.067		-0.055		-0.121		-0.005
			(-4.15)		(-2.72)		(-4.06)		(-0.09)
$\delta_{(P)}$	1.418	-1.630	-2.133	-9.393	-10.056	-10.629	-12.959	-3.340	-3.199
	(1.99)	(-0.84)	(-1.02)	(-3.47)	(-3.60)	(-2.79)	(-2.95)	(-1.55)	(-1.42)
δ_N		3.582	3.902	6.958	7.229	12.811	14.649	6.090	6.314
		(2.05)	(2.05)	(3.30)	(3.24)	(4.39)	(4.85)	(3.82)	(3.65)
$\phi(1)$	0.139	0.171	0.130	0.227	0.185	0.240	0.208	0.190	0.191
	(5.92)	(5.81)	(4.31)	(7.10)	(4.89)	(6.50)	(5.61)	(6.41)	(5.93)
$AdjR^2(\%)$	1.79	1.88	2.18	2.25	2.39	2.61	2.79	2.10	2.08
Panel B. 2nd Sub-period: Apr. 2, 1951 – Dec. 31, 2019									
$c_{(P)}(\times 100)$	0.011	0.011	0.095	0.015	0.106	0.025	0.089	0.017	0.032
	(1.15)	(1.07)	(4.91)	(1.18)	(4.42)	(3.23)	(5.07)	(2.39)	(0.67)
$c_N(\times 100)$			-0.070		-0.100		-0.091		-0.064
			(-4.78)		(-3.74)		(-3.30)		(-1.33)
$\delta_{(P)}$	2.069	-1.223	-2.510	-3.554	-4.843	-7.910	-10.040	-6.277	-5.865
	(1.83)	(-0.44)	(-1.06)	(-1.12)	(-1.54)	(-3.04)	(-2.93)	(-2.28)	(-1.85)
δ_N		4.620	5.129	7.393	8.480	6.840	9.450	9.459	10.441
		(2.43)	(2.41)	(2.09)	(2.13)	(2.20)	(2.45)	(4.85)	(4.52)
$\phi(1)$	0.041	0.073	0.018	0.100	0.034	0.089	0.059	0.107	0.098
	(1.47)	(2.84)	(0.59)	(3.86)	(1.05)	(2.34)	(1.48)	(4.03)	(3.15)
$AdjR^2(\%)$	0.58	0.66	1.08	0.76	1.08	0.70	0.79	1.11	1.09

Note: This table reports the estimates of the following models for the value-weighted market portfolio for two sub-periods.

Model 1: $r_{m,t+1} = c + \delta\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 2: $r_{m,t+1} = c + (\delta_P Pd + \delta_N Nd)\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 3: $r_{m,t+1} = (c_P Pd + c_N Nd) + (\delta_P Pd + \delta_N Nd)\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

where $r_{m,t+1}$ is daily realized excess return of the value-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance. The RRA parameter is measured by $\delta_P(\delta_N)$ under prior *d*-day positive (negative) returns.

Table 5

Estimation Results of the Indirect Relationship between Excess Market Returns and Contemporary Volatility Innovations

	Full-period: Jul. 1965 – Dec. 2019				1st Sub-period: Jul. 1965 – Dec. 1987				2nd Sub-period: Apr. 1951 – Dec. 2019			
	[A]	[B]	[C]	[D]	[A]	[B]	[C]	[D]	[A]	[B]	[C]	[D]
$c_{(P)}(\times 100)$	0.030	0.094	0.009	0.079	0.025	0.097	0.007	0.084	0.032	0.110	0.008	0.095
	(4.59)	(9.13)	(1.24)	(6.77)	(3.17)	(7.84)	(0.89)	(6.19)	(4.63)	(9.00)	(0.87)	(6.70)
$c_N(\times 100)$		-0.037		-0.065		-0.051		-0.073		-0.049		-0.076
		(-2.81)		(-5.23)		(-3.26)		(-5.22)		(-3.30)		(-4.65)
π_P	3.179	3.166	3.193	3.179	4.507	4.509	4.501	4.502	-1.558	-1.716	-1.396	-1.544
	(2.08)	(2.06)	(2.11)	(2.08)	(3.51)	(3.48)	(3.52)	(3.49)	(-0.54)	(-1.54)	(-0.50)	(-0.56)
π_N	-3.150	-3.173	-3.303	-3.339	-3.365	-3.401	-3.427	-3.468	-4.610	-4.619	-4.784	-4.810
	(-2.36)	(-2.38)	(-2.58)	(-2.63)	(-2.15)	(-2.18)	(-2.24)	(-2.29)	(-3.23)	(-3.20)	(-3.47)	(-3.53)
δ_P			-1.532	-2.164			-1.469	-2.001			-1.748	-3.085
			(-0.81)	(-1.08)			(-0.59)	(-0.73)			(-0.72)	(-1.37)
δ_N			4.836	5.247			4.144	4.491			6.511	7.041
			(3.54)	(3.60)			(2.90)	(2.88)			(3.29)	(3.36)
$\phi(1)$	0.040	-0.001	0.096	0.057	0.118	0.072	0.167	0.123	-0.007	-0.069	0.059	0.000
	(1.56)	(-0.04)	(3.59)	(1.99)	(4.53)	(2.61)	(5.24)	(3.63)	(-0.20)	(-1.77)	(1.85)	(0.00)
$AdjR^2(\%)$	2.44	2.66	2.67	2.93	4.92	5.21	5.10	5.43	2.90	3.31	3.19	3.65

Note: This table reports the estimation results of the indirect risk-return relation in the generalized specification of Model 4 for the value-weighted market portfolio for the full period (Jan. 2, 1926 – Dec. 31, 2019) and two sub-periods (Jan. 2, 1926 – Dec. 31, 1987 and Apr. 2, 1951 – Dec. 31, 2019).

Model 4: $r_{m,t+1} = (c_P P + c_N N) + (\pi_P P + \pi_N N)\hat{\eta}^2_{m,t+1} + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

where $r_{m,t+1}$ is daily realized excess return of the value-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional. $P = 1$ when $e_{m,t} > 0$ and $N = 1$ when $e_{m,t} < 0$, where $e_{m,t}$ is the mean-deviated excess market returns. $\hat{\eta}^2_{m,t+1}$ is the contemporaneous volatility innovation that represents unexpected volatility changes (i.e., $\hat{\eta}^2_{m,t+1} = e^2_{m,t+1} - \hat{\sigma}^2_{m,t+1}$). The indirect risk-return relation is measured by $\pi_P(\pi_N)$ under a prior positive (negative) return.

Table 6

Estimation Results of Asymmetric Intertemporal Risk-Return Relation under High/Low Investor Sentiment

	Constant RRA under High/Low Market Sentiment		Prior 1-day Positive/Negative Price Changes		Prior 2-day Positive/Negative Price Changes		Prior 3-day Positive/Negative Price Changes		One Standard Dev. Positive/Negative Return Changes		Two Standard Dev. Positive/Negative Return Changes	
$c^{(H)}(\times 100)$	0.007	-0.017	0.007	-0.017	0.011	0.018	0.022	0.006	0.013	-0.090	0.019	-0.108
	(0.60)	(-1.21)	(0.59)	(-1.16)	(0.86)	(1.02)	(2.30)	(0.49)	(1.52)	(-1.65)	(2.18)	(-0.76)
$c^{L}(\times 100)$		0.025		0.024		0.008		0.035		0.060		0.058
		(1.44)		(1.54)		(0.42)		(2.56)		(1.22)		(0.51)
$\delta^{H}_{(P)}$	1.929	2.946	1.666	2.695	-6.562	-6.866	-11.486	-10.494	-6.537	-6.596	-7.110	-7.478
	(1.96)	(2.33)	(1.04)	(1.69)	(-2.10)	(-2.07)	(-3.40)	(-3.06)	(-2.85)	(-3.37)	(-2.02)	(-2.05)
δ^{L}_{P}			-3.236	-3.737	-3.280	-3.182	-6.679	-7.123	-4.296	-4.038	-4.112	-4.329
			(-1.85)	(-2.37)	(-0.82)	(-0.76)	(-2.09)	(-2.17)	(-1.97)	(-2.09)	(-0.99)	(-1.25)
$\delta^{H}_{(N)}$	1.510	0.980	2.179	3.145	2.876	2.645	8.568	9.014	7.207	8.217	7.096	8.288
	(0.92)	(0.58)	(1.20)	(1.49)	(0.78)	(0.70)	(2.71)	(2.71)	(3.80)	(3.24)	(2.03)	(2.07)
δ^{L}_{N}			5.407	4.873	10.040	10.143	14.379	13.832	8.020	8.704	8.878	8.968
			(2.20)	(2.01)	(3.79)	(3.75)	(4.77)	(4.68)	(3.16)	(3.05)	(2.80)	(2.33)
$\phi_H(1)$	0.076	0.086	0.079	0.088	0.103	0.102	0.145	0.147	0.123	0.102	0.094	0.091
	(2.15)	(2.39)	(2.14)	(2.36)	(2.56)	(2.49)	(3.81)	(3.84)	(3.29)	(2.41)	(2.91)	(2.92)
$\phi_L(1)$	-0.014	-0.020	0.046	0.040	0.089	0.090	0.027	0.023	0.060	0.040	0.021	0.019
	(-0.29)	(-0.41)	(1.13)	(0.98)	(2.62)	(2.61)	(0.55)	(0.46)	(1.30)	(0.82)	(0.46)	(0.47)
$AdjR^2(\%)$	0.59	0.61	0.73	0.75	0.96	0.95	0.76	0.78	1.02	1.04	0.88	0.86

Note: This table reports estimates of the constant (Model 5) and asymmetric (Model 6) intertemporal risk-return relation under high/low sentiment for the value-weighted market portfolio for the period over July 1965 – December 2018:

Model 5: Constant RRA under high/low sentiment

$$r_{m,t+1} = \left[c^H + \delta^H \hat{\sigma}^2_{m,t+1} + \phi_H(1) r_{m,t}\right] \cdot H + \left[c^L + \delta^L \hat{\sigma}^2_{m,t+1} + \phi_L(1) r_{m,t}\right] \cdot L + \varepsilon_{m,t+1},$$

Model 6: Asymmetric RRA under high/low sentiment

$$r_{m,t+1} = \left[c^H + (\delta^H_P Pd + \delta^H_N Nd)\hat{\sigma}^2_{m,t+1} + \phi_H(1) r_{m,t}\right] \cdot H + \left[c^L + (\delta^L_P Pd + \delta^L_N Nd)\hat{\sigma}^2_{m,t+1} + \phi_L(1) r_{m,t}\right] \cdot L + \varepsilon_{m,t+1},$$

where $r_{m,t+1}$ is daily realized excess return of the value-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance. $H(L)$ is the dummy representing high-(low-) sentiment regimes. $Pd(Nd)$ is the dummy to capture not only prior d-day positive (negative) returns but also prior one and two standard deviation of positive and negative return changes. The RRA parameter in the high-sentiment regime is measured by $\delta^H_P(\delta^H_N)$ under prior positive (negative) returns, while $\delta^L_P(\delta^L_N)$ measures the RRA parameter under prior positive (negative) returns in the low-sentiment regime. The price adjustment during the high-sentiment regime is measured by $\phi_H(1)$, while it is measured by $\phi_L(1)$ during the low-sentiment regime.

Table 7

Robustness Check of Asymmetric Intertemporal Risk-Return Relation to the Alternative Proxies for Investor Sentiment

Panel A. University of Michigan Consumer Sentiment Index

	Prior 1-day Positive/Negative Price Changes		Prior 2-day Positive/Negative Price Changes		Prior 3-day Positive/Negative Price Changes		One Standard Dev. Positive/Negative Return Changes		Two Standard Dev. Positive/Negative Return Changes	
$c^{(H)}(\times 100)$	0.009	0.014	0.007	-0.007	0.021	0.019	0.014	0.017	0.018	0.019
	(0.66)	(0.71)	(0.45)	(-0.39)	(2.33)	(1.89)	(1.63)	(1.62)	(2.27)	(1.88)
$c^{L}(\times 100)$		0.002		0.022		0.023		0.009		0.017
		(0.11)		(1.13)		(1.49)		(0.68)		(1.33)
δ_P^H	-0.366	-0.642	1.307	2.212	-2.799	-2.682	-6.383	-6.473	-5.248	-5.258
	(-0.14)	(-0.23)	(0.29)	(0.50)	(-0.79)	(-0.78)	(-2.81)	(-2.84)	(-1.76)	(-1.78)
δ_P^L	-2.060	-1.863	-5.922	-6.415	-10.128	-10.197	-4.603	-4.522	-4.648	-4.633
	(-1.02)	(-0.89)	(-2.40)	(-2.40)	(-4.04)	(-3.91)	(-2.04)	(-1.98)	(-1.30)	(-1.29)
δ_N^H	2.785	2.604	9.592	10.006	4.088	4.151	7.805	7.747	8.338	8.331
	(0.85)	(0.77)	(2.08)	(1.94)	(0.80)	(0.80)	(2.97)	(2.95)	(2.09)	(2.10)
δ_N^L	4.962	5.150	5.789	5.362	7.392	7.350	8.054	8.131	7.528	7.542
	(2.07)	(2.08)	(2.03)	(1.99)	(2.04)	(2.01)	(3.59)	(3.56)	(3.34)	(3.33)
$\phi(1)$	0.056	0.057	0.092	0.092	0.072	0.072	0.088	0.088	0.059	0.059
	(2.09)	(2.12)	(3.09)	(3.11)	(2.13)	(2.13)	(2.82)	(2.83)	(1.93)	(1.94)
$AdjR^2(\%)$	0.555	0.550	0.781	0.791	0.623	0.615	0.887	0.880	0.852	0.844

Panel B. Augmented Sentiment Index

$c^{(H)}(\times 100)$	-0.006	-0.008	0.007	0.037	0.017	0.038	0.008	0.021	0.015	0.032
	(-0.57)	(-0.62)	(0.44)	(2.24)	(1.75)	(3.05)	(0.88)	(1.79)	(1.81)	(2.91)
$c^{L}(\times 100)$		-0.005		-0.022		-0.004		-0.006		-0.003
		(-0.28)		(-0.95)		(-0.29)		(-0.50)		(-0.26)
δ_P^H	0.910	1.012	-6.956	-8.713	-11.874	-13.609	-3.756	-4.135	-0.314	-0.600
	(0.33)	(0.36)	(-1.63)	(-1.94)	(-2.47)	(-2.67)	(-1.45)	(-1.61)	(-0.13)	(-0.24)
δ_P^L	-2.737	-2.789	-3.568	-2.548	-7.196	-6.398	-6.460	-6.200	-6.988	-6.757
	(-1.24)	(-1.31)	(-1.15)	(-0.78)	(-2.42)	(-1.96)	(-2.32)	(-2.13)	(-1.95)	(-1.79)
δ_N^H	10.879	10.975	11.803	10.053	15.746	14.504	17.398	17.022	18.854	18.536
	(3.74)	(3.75)	(4.08)	(3.60)	(4.41)	(4.08)	(6.61)	(6.70)	(6.41)	(6.11)
δ_N^L	2.750	2.713	6.926	7.444	4.609	4.995	5.487	5.667	4.992	5.144

	(1.71)	(1.68)	(2.35)	(2.53)	(1.66)	(1.73)	(3.28)	(2.77)	(2.79)	(2.36)
$\phi(1)$	0.065	0.065	0.108	0.104	0.092	0.089	0.092	0.091	0.059	0.057
	(2.37)	(2.39)	(3.73)	(3.65)	(2.72)	(2.64)	(2.51)	(2.49)	(1.65)	(1.60)
$AdjR^2(\%)$	0.802	0.795	0.760	0.822	0.724	0.756	1.249	1.258	1.227	1.252

Note: This table reports the estimation results of the robustness check of the asymmetric intertemporal risk-return relation to the alternative proxies for sentiment for the period over July 1965 – December 2018 in the following model:

$$r_{m,t+1} = \left[c^H + (\delta_P^H Pd + \delta_N^H Nd)\hat{\sigma}_{m,t+1}^2\right] \cdot H + \left[c^L + (\delta_P^L Pd + \delta_N^L Nd)\hat{\sigma}_{m,t+1}^2\right] \cdot L + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1},$$

where $r_{m,t+1}$ is daily realized excess return of the value-weighted market portfolio, and $\hat{\sigma}_{m,t+1}^2$ is its conditional variance. $H(L)$ is the dummy representing high- (low-) sentiment regimes. As for the alternative measures of investor sentiment, we employ the University of Michigan Consumer Sentiment Index and the Augmented Sentiment Index that used by Doukas and Han (2021). $Pd(Nd)$ is the dummy to capture not only prior d-day positive (negative) returns but also prior one and two standard deviation of positive and negative return changes. The RRA parameter in the high-sentiment regime is measured by $\delta_P^H(\delta_N^H)$ under prior positive (negative) returns, while $\delta_P^L(\delta_N^L)$ measures the RRA parameter under prior positive (negative) returns in the low-sentiment regime.

Table 8

Estimation Results of the Asymmetric Intertemporal Risk-Return Relation Conditioned on Business Cycle

	Constant RRA under Expansion/Recession Business Cycle	Prior 1-day Positive/Negative Price Changes		Prior 2-day Positive/Negative Price Changes		Prior 3-day Positive/Negative Price Changes		One Standard Dev. Positive/Negative Return Changes		Two Standard Dev. Positive/Negative Return Changes	
$\mu^E_{(P)}(\times 100)$	0.005	0.003	0.072	0.030	0.090	0.030	0.093	0.027	0.018	0.032	0.005
	(0.60)	(0.37)	(5.31)	(2.92)	(4.35)	(4.02)	(5.37)	(4.30)	(0.51)	(5.29)	(0.04)
$\mu^R_P(\times 100)$			0.068		0.071		0.052		-0.022		-0.115
			(1.99)		(1.72)		(1.17)		(-0.21)		(-0.58)
$\mu^E_N(\times 100)$			-0.056		-0.045		-0.081		-0.013		0.103
			(-3.14)		(-1.53)		(-1.97)		(-0.25)		(1.12)
$\mu^R_{(N)}(\times 100)$	-0.049	-0.046	-0.163	-0.058	-0.211	-0.054	-0.308	-0.057	-0.139	-0.047	-0.146
	(-2.00)	(-1.74)	(-3.91)	(-1.82)	(-3.48)	(-1.79)	(-3.71)	(-2.23)	(-1.41)	(-1.88)	(-0.77)
$\delta^E_{(P)}$	4.162	2.757	1.288	-5.445	-6.705	-8.829	-11.852	-3.296	-2.621	-3.014	-2.673
	(3.73)	(1.60)	(0.77)	(-1.89)	(-2.41)	(-3.47)	(-4.17)	(-1.66)	(-1.10)	(-1.57)	(-0.58)
δ^R_P		-2.492	-3.333	-6.666	-7.696	-8.598	-10.127	-2.153	-2.349	-0.735	-0.186
		(-1.08)	(-1.48)	(-2.45)	(-2.58)	(-2.43)	(-2.59)	(-0.99)	(-0.88)	(-0.32)	(-0.06)
δ^E_N		5.678	6.527	6.980	8.162	11.503	14.395	8.999	9.616	9.568	8.734
		(3.15)	(2.96)	(2.10)	(1.99)	(3.37)	(2.77)	(4.44)	(3.92)	(3.08)	(2.83)
$\delta^R_{(N)}$	1.121	3.583	4.346	8.751	9.811	11.706	14.998	5.942	6.718	5.130	6.246
	(1.19)	(2.07)	(1.97)	(3.78)	(3.38)	(3.31)	(3.60)	(4.10)	(3.10)	(2.74)	(2.06)
$\phi(1)$	0.072	0.103	0.067	0.151	0.111	0.153	0.127	0.122	0.120	0.100	0.109
	(3.12)	(4.07)	(2.54)	(5.47)	(3.56)	(5.01)	(3.91)	(4.73)	(4.32)	(4.22)	(4.47)
$AdjR^2(\%)$	0.89	1.00	1.24	1.23	1.36	1.38	1.49	1.15	1.07	1.10	1.01

Note: This table reports the estimates of the constant (Model 7) and asymmetric (Model 8) intertemporal risk-return relation for the value-weighted market portfolio conditioned on expansion and recession periods for the full period Jan. 2, 1926 – Dec. 31, 2019:

Model 7: Constant RRA under expansion/recession period

$$r_{m,t+1} = \left[\mu^E + \delta^E \hat{\sigma}^2_{m,t+1}\right] \cdot E + \left[\mu^R + \delta^R \hat{\sigma}^2_{m,t+1}\right] \cdot R + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1},$$

Model 8: Asymmetric RRA under expansion/recession in business cycle

$$r_{m,t+1} = \left[(\mu^E_P Pd + \mu^E_N Nd) + (\delta^E_P Pd + \delta^E_N Nd)\hat{\sigma}^2_{m,t+1}\right] \cdot E + \left[(\mu^R_P Pd + \mu^R_N Nd) + (\delta^R_P Pd + \delta^R_N Nd)\hat{\sigma}^2_{m,t+1}\right] \cdot R + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1},$$

where $r_{m,t+1}$ is daily realized excess return of the value-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance. $Pd(Nd)$ is the dummy to capture not only prior d-day positive (negative) returns but also prior one and two standard deviation positive and negative returns. $E(R)$ is the dummy variable representing the expansion (recession) period in the business cycle defined by the NBER. Periods of expansion begin at the trough date and end at the peak date, while periods of recession begin at the peak date and end at the trough date. The RRA parameter in the expansion periods is measured by $\delta^E_P\,(\delta^E_N)$ under prior positive (negative) returns, while $\delta^R_P\,(\delta^R_N)$ measures the RRA parameter under prior positive (negative) returns in the recession periods in business cycle.

Table 9
Estimation Results of the Short-sale Effect of the Introduction of Options on the Asymmetric Intertemporal Risk-Return Relation

	1st Case of P and N: Prior 1-day positive-negative price change			2nd Case of P and N: One standard deviation return change			3rd Case of P and N: Two standard deviations return change		
	Model 9A	Model 9B	Model 9C	Model 9A	Model 9B	Model 9C	Model 9A	Model 9B	Model 9C
$c_{(1)}(\times 100)$	0.012	0.014	0.081	0.015	0.014	0.015	0.019	0.018	-0.061
	(1.58)	(1.53)	(6.89)	(2.21)	(1.47)	(0.35)	(2.84)	(1.89)	(-0.70)
$c_2\,(\times 100)$		-0.004	-0.059		0.003	-0.028		0.003	0.069
		(-0.26)	(-4.36)		(0.21)	(-0.55)		(0.25)	(0.67)
δ_P	0.211	0.160	-0.330	-0.392	-0.371	-0.144	1.190	1.208	1.888
	(0.11)	(0.08)	(-0.17)	(-0.18)	(-0.18)	(-0.07)	(0.53)	(0.54)	(0.86)
δ_N	1.887	1.849	2.264	4.151	4.168	4.630	3.818	3.833	3.556
	(1.14)	(1.13)	(1.20)	(2.68)	(2.84)	(2.96)	(1.80)	(1.81)	(1.50)
δ_P^O	-3.945	-3.814	-4.115	-5.776	-5.830	-5.778	-6.902	-6.945	-6.903
	(-2.03)	(-1.96)	(-2.01)	(-2.10)	(-2.30)	(-2.13)	(-2.25)	(-2.27)	(-2.30)
δ_N^O	4.600	4.702	4.620	4.489	4.444	4.493	4.995	4.958	5.031
	(2.06)	(1.99)	(1.99)	(2.07)	(2.04)	(2.09)	(1.94)	(1.93)	(1.98)
$\phi(1)$	0.098	0.098	0.060	0.114	0.114	0.112	0.094	0.094	0.104
	(4.30)	(4.09)	(2.36)	(4.91)	(4.91)	(3.84)	(4.15)	(4.14)	(4.30)
$AdjR^2(\%)$	0.98	0.97	1.23	1.14	1.14	1.12	1.11	1.10	1.08
$\delta_P+\delta_P^O$	-3.734	-3.655	-4.445	-6.168	-6.201	-5.922	-5.712	-5.737	-5.015
(*t*-value)	(-2.69)	(-2.28)	(-2.94)	(-3.37)	(-3.34)	(-3.27)	(-2.56)	(-2.57)	(-2.05)
$\delta_N+\delta_N^O$	6.486	6.551	6.884	8.639	8.612	9.123	8.813	8.791	8.586
(*t*-value)	(3.10)	(2.82)	(2.82)	(5.08)	(5.02)	(4.93)	(3.56)	(3.59)	(3.47)
F-value	19.042	19.033	19.761	18.272	18.260	18.282	18.008	17.996	18.118
LR-value	38.084	38.066	39.522	36.544	36.520	36.563	36.016	35.992	36.236

Note: This table reports estimates that measure the effect of the introduction of options trading on the intertemporal risk-return relation. The following three models are estimated for the value-weighted market portfolio for the full period Jan. 2, 1926 – Dec. 31, 2019:

Model 9A: $r_{m,t+1} = c + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta_P^O P + \delta_N^O N)\hat{\sigma}^2_{m,t+1} \cdot O + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 9B: $r_{m,t+1} = (c_1 + c_2 O) + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta_P^O P + \delta_N^O N)\hat{\sigma}^2_{m,t+1} \cdot O + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 9C: $r_{m,t+1} = (c_1 P + c_2 N) + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta_P^O P + \delta_N^O N)\hat{\sigma}^2_{m,t+1} \cdot O + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

where $r_{m,t+1}$ is daily realized excess return of the value-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance. O is the dummy variable to represent the period after the introduction of options into the markets (January 1981 – December 2019). $P(N)$ is the dummy to represent three different cases of state-dependency. For the first case, $P(N)$ captures prior 1-day positive (negative) return. For the second and third cases, $P(N)$ captures prior one and two standard deviation positive (negative) returns, respectively. $\delta_P^O\,(\delta_N^O)$ measures the differential effect of the introduction of options on the intertemporal risk-return relation.

$\delta_P + \delta_P^O (\delta_N + \delta_N^O)$ is the RRA parameter after the introduction of options. We perform three diagnostic tests on the importance of the introduction of options. We report the t-value for $H_0\colon \delta_P + \delta_P^O = 0$ and $H_0\colon \delta_N + \delta_N^O = 0$, respectively, while we report the F-value and the LR-value on the joint nulls of $\delta_P^O = 0$ and $\delta_N^O = 0$.

Table 10

Results of the Placebo Test on the Effect of Introduction of Options on the Asymmetric Intertemporal Risk-Return Relation

	Panel A. 10-years placebo period before the introduction of options						Panel B. 20-years placebo period before the introduction of options					
	Prior 1-day positive-negative price change			One standard deviation return change			Prior 1-day positive-negative price change			One standard deviation return change		
	[A]	[B]	[C]	[A]	[B]	[C]	[A]	[B]	[C]	[A]	[B]	[C]
$c_{(1)}(\times 100)$	0.014	0.014	0.074	0.016	0.016	-0.004	0.014	0.016	0.067	0.016	0.017	-0.013
	(1.86)	(1.79)	(6.00)	(2.37)	(2.30)	(-0.09)	(1.84)	(1.91)	(5.37)	(2.33)	(2.26)	(-0.28)
$c_2\,(\times 100)$		-0.001	-0.049		-0.004	0.009		-0.015	-0.042		-0.008	0.013
		(-0.03)	(-3.33)		(-0.20)	(0.19)		(-0.85)	(-2.83)		(-0.52)	(0.26)
δ_P	-1.537	-1.538	-2.050	-2.715	-2.722	-2.493	-1.688	-1.751	-2.124	-2.817	-2.843	-2.528
	(-0.94)	(-0.94)	(-1.19)	(-1.48)	(-1.49)	(-1.29)	(-1.02)	(-1.06)	(-1.23)	(-1.52)	(-1.53)	(-1.30)
δ_N	3.735	3.733	4.061	6.180	6.174	6.420	3.787	3.736	4.068	6.172	6.147	6.424
	(2.18)	(2.17)	(2.22)	(4.25)	(4.24)	(4.08)	(2.18)	(2.14)	(2.21)	(4.20)	(4.19)	(4.06)
δ_P^O	14.250	14.311	11.929	14.083	14.293	14.671	17.102	18.232	14.888	15.314	15.648	15.998
	(3.81)	(3.05)	(3.21)	(3.32)	(3.54)	(3.42)	(6.06)	(5.51)	(5.31)	(4.50)	(4.65)	(4.63)
δ_N^O	-14.197	-14.149	-12.143	-14.830	-14.663	-14.523	-14.515	-13.622	-12.566	-11.599	-11.318	-11.357
	(-6.34)	(-5.33)	(-5.37)	(-3.59)	(-3.31)	(-3.46)	(-7.38)	(-6.01)	(-6.44)	(-3.43)	(-3.19)	(-3.15)
$\phi(1)$	0.089	0.089	0.057	0.110	0.110	0.116	0.080	0.080	0.054	0.108	0.108	0.116
	(3.62)	(3.62)	(2.24)	(4.30)	(4.30)	(4.13)	(3.26)	(3.26)	(2.09)	(4.20)	(4.19)	(4.11)
$AdjR^2(\%)$	1.02	1.01	1.20	1.11	1.11	1.09	1.19	1.19	1.33	1.14	1.14	1.12

Note: This table reports the results of the Placebo test on the effect of introduction of options on the asymmetric intertemporal risk-return relation. The following three models are estimated for the value-weighted market portfolio for the full period Jan. 2, 1926 – Dec. 31, 2019:

Model 10A: $r_{m,t+1} = c + (\delta_P P + \delta_N N)\hat{\sigma}_{m,t+1}^2 + (\delta_P^{CB} P + \delta_N^{CB} N)\hat{\sigma}_{m,t+1}^2 \cdot CB + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1},$

Model 10B: $r_{m,t+1} = (c_1 + c_2 CB) + (\delta_P P + \delta_N N)\hat{\sigma}_{m,t+1}^2 + (\delta_P^{CB} P + \delta_N^{CB} N)\hat{\sigma}_{m,t+1}^2 \cdot CB + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1},$

Model 10C: $r_{m,t+1} = (c_1 P + c_2 N) + (\delta_P P + \delta_N N)\hat{\sigma}_{m,t+1}^2 + (\delta_P^{CB} P + \delta_N^{CB} N)\hat{\sigma}_{m,t+1}^2 \cdot CB + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1},$

where $r_{m,t+1}$ is daily realized excess return of the value-weighted market portfolio, and $\hat{\sigma}_{m,t+1}^2$ is its conditional variance. CB is the dummy variable to represent the three randomized placebo periods before the introduction of options into the markets. $P(N)$ captures prior 1-day positive (negative) return and prior one standard deviation positive (negative) returns, respectively. δ_P^{CB} and δ_N^{CB} measure the placebo effect on the intertemporal risk-return relation.

Table A.1
Estimation Results of the Intertemporal Risk-Return Relation for the Equal-Weighted Market Portfolio

	Constant RRA Parameter	Prior 1-day Positive/Negative Price Changes		Prior 2-day Consecutive Positive/Negative Price Changes		Prior 3-day Consecutive Positive/Negative Price Changes		One Standard Deviation Positive/Negative Return Changes	
	Model 1	Model 2	Model 3	Model 2	Model 3	Model 2	Model 3	Model 2	Model 3
$c_{(P)}(\times 100)$	0.013	0.012	0.051	0.014	0.053	0.026	0.053	0.025	0.053
	(1.83)	(1.68)	(5.09)	(1.60)	(3.27)	(3.34)	(3.25)	(3.74)	(1.40)
$c_N(\times 100)$			-0.039		-0.088		-0.131		0.005
			(-2.70)		(-3.82)		(-3.77)		(0.11)
$\delta_{(P)}$	3.565	1.112	1.234	-0.362	0.269	-1.562	-1.024	-0.397	-0.638
	(5.58)	(0.87)	(0.94)	(-0.28)	(0.21)	(-0.73)	(-0.47)	(-0.26)	(-0.41)
δ_N		5.725	5.787	10.377	10.951	12.066	14.151	7.613	7.782
		(3.80)	(3.73)	(5.15)	(5.34)	(5.53)	(5.86)	(4.27)	(4.17)
$\phi(1)$	0.303	0.337	0.309	0.378	0.338	0.371	0.343	0.355	0.357
	(11.49)	(10.81)	(9.24)	(11.98)	(9.30)	(13.12)	(10.10)	(12.02)	(11.90)
$AdjR^2(\%)$	5.69	5.78	5.88	6.20	6.33	6.08	6.17	5.99	5.95

Note: This table reports the estimates of the following models for the equal-weighted market portfolio for the full period Jan. 2, 1926 – Dec. 31, 2019:

Model 1: $r_{m,t+1} = c + \delta\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 2: $r_{m,t+1} = c + (\delta_P Pd + \delta_N Nd)\cdot\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 3: $r_{m,t+1} = (c_P Pd + c_N Nd) + (\delta_P Pd + \delta_N Nd)\cdot\hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5}\phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

where $r_{m,t+1}$ is daily realized excess return of the equal-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance estimated from the EGARCH (1, 2) model. ϕ_j is the j^{th} order return autocorrelation coefficient. $\phi(1)$ is the sum of autocorrelation coefficients, i.e., $\phi(1) = \sum_{j=1}^{5}\phi_j$. Pd (Nd) is the dummy to capture prior *d*-day positive (negative) returns, such that $P2 = 1$ when $e_{m,t-1} > 0$ and $e_{m,t} > 0$ $(N2 = 1$ when $e_{m,t-1} < 0$ and $e_{m,t} < 0)$ where $e_{m,t}$ is the mean-deviated excess market returns. The RRA parameter is measured by δ_P (δ_N) under prior *d*-day positive (negative) returns. We also estimate δ_P and δ_N using one standard deviation positive/negative return changes as the dummy variables. The numbers in parentheses are the Newey-West (1987) adjusted *t*-statistics. *Adj.*R^2 (%) is the percentage adjusted R^2.

Table A.2
Estimation Results of the Intertemporal Risk-Return Relation under Prior Extreme Price Changes for the Equal-Weighted Market Portfolio

	Prior 4-day Positive/Negative Price Changes		Prior 5-day Consecutive Positive/Negative Price Changes		Prior 6-day Consecutive Positive/Negative Price Changes		Two Standard Deviation Positive/Negative Return Changes	
	Model 2	Model 3	Model 2	Model 3	Model 2	Model 3	Model 2	Model 3
$c_{(P)}(\times 100)$	0.036	0.057	0.037	0.069	0.041	0.084	0.032	-0.087
	(5.03)	(2.90)	(5.30)	(2.70)	(6.17)	(3.98)	(4.96)	(-0.54)
$c_N(\times 100)$		-0.169		-0.261		-0.260		0.158
		(-3.06)		(-4.28)		(-3.40)		(1.40)
$\delta_{(P)}$	-0.784	-0.650	-2.767	-3.274	2.363	0.719	-6.033	-5.085
	(-0.35)	(-0.26)	(-0.67)	(-0.72)	(0.65)	(0.19)	(-2.73)	(-1.55)
δ_N	11.147	14.878	17.603	23.336	16.904	23.438	8.306	7.629
	(3.63)	(3.80)	(5.82)	(6.50)	(5.97)	(6.52)	(3.68)	(3.19)
$\phi(1)$	0.339	0.323	0.351	0.338	0.312	0.309	0.363	0.383
	(13.31)	(11.36)	(10.34)	(9.50)	(10.08)	(9.54)	(13.86)	(15.21)
$AdjR^2(\%)$	5.70	5.72	5.99	6.04	5.68	5.64	5.45	5.33

Note: This table reports the estimates of following models for the equal-weighted market portfolio for the full period Jan. 2, 1926 – Dec. 31, 2019:

Model 2: $r_{m,t+1} = c + (\delta_P Pd + \delta_N Nd) \cdot \hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 3: $r_{m,t+1} = (c_P Pd + c_N Nd) + (\delta_P Pd + \delta_N Nd) \cdot \hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

where $r_{m,t+1}$ is daily realized excess return of the equal-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance. Pd (Nd) is the dummy to capture prior *extreme* d-day positive (negative) returns, i.e., d = 4, 5, 6. We also estimate the models using two standard deviation positive/negative return changes as the dummy variables. The RRA parameter is measured by δ_P (δ_N) under prior *extreme* positive (negative) price change(s).

Table A.3
Estimation Results of the Indirect Risk-Return Relationship for the Equal-Weighted Market Portfolio

	Full-period: Jul. 1965 – Dec. 2019				1st Sub-period: Jul. 1965 – Dec. 1987				2nd Sub-period: Apr. 1951 – Dec. 2019			
	[A]	[B]	[C]	[D]	[A]	[B]	[C]	[D]	[A]	[B]	[C]	[D]
$c_{(P)}(\times 100)$	0.056	0.096	0.010	0.053	0.053	0.099	0.008	0.057	0.026	0.076	0.023	0.074
	(8.85)	(9.66)	(1.26)	(4.89)	(6.77)	(8.34)	(0.80)	(4.27)	(3.15)	(3.27)	(1.60)	(2.92)
$c_N(\times 100)$		0.006		-0.048		-0.004		-0.058		-0.035		-0.038
		(0.39)		(-3.25)		(-0.19)		(-3.09)		(-1.38)		(-1.26)
π_P	3.888	3.887	3.926	3.928	4.361	4.370	4.357	4.372	2.466	2.496	4.601	4.500
	(1.81)	(1.81)	(1.82)	(1.82)	(1.80)	(1.81)	(1.76)	(1.77)	(0.50)	(0.51)	(0.76)	(0.74)
π_N	-4.356	-4.382	-4.636	-4.669	-4.045	-4.083	-4.314	-4.362	-8.697	-8.778	-9.216	-9.358
	(-3.91)	(-3.96)	(-4.39)	(-4.45)	(-3.48)	(-3.55)	(-4.07)	(-4.17)	(-5.27)	(-5.26)	(-4.91)	(-5.09)
δ_P			1.559	1.695			1.448	1.643			5.678	5.340
			(1.02)	(1.09)			(0.98)	(1.09)			(1.15)	(1.11)
δ_N			6.847	6.925			6.157	6.224			-2.689	-3.010
			(4.25)	(4.22)			(4.13)	(4.16)			(-1.00)	(-1.13)
$\phi(1)$	0.256	0.228	0.318	0.286	0.289	0.259	0.343	0.310	0.425	0.366	0.390	0.328
	(8.50)	(6.83)	(9.45)	(7.76)	(7.84)	(6.47)	(8.75)	(7.17)	(11.35)	(6.54)	(13.21)	(6.28)
$AdjR^2(\%)$	8.01	8.11	8.74	8.88	10.76	10.89	11.46	11.63	20.20	20.53	20.50	20.84

Note: This table reports the estimates of the indirect risk-return relation in the generalized specification of Model 3 for the equal-weighted market portfolio for the full period (Jan. 2, 1926 – Dec. 31, 2019) and two sub-periods (Jan. 2, 1926 – Dec. 31, 1987 and Apr. 2, 1951 – Dec. 31, 2019).

Model 4: $r_{m,t+1} = (c_P P + c_N N) + (\pi_P P + \pi_N N) \cdot \hat{\eta}^2_{m,t+1} + (\delta_P P + \delta_N N) \cdot \hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

where $r_{m,t+1}$ is daily realized excess return of the equal-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance. $P = 1$ when $e_{m,t} > 0$ and $N = 1$ when $e_{m,t} < 0$, where $e_{m,t}$ is the mean-deviated excess market returns. $\hat{\eta}^2_{m,t+1}$ is the contemporaneous volatility innovation that represents the unexpected volatility (i.e., $\hat{\eta}^2_{m,t+1} = e^2_{m,t+1} - \hat{\sigma}^2_{m,t+1}$). The indirect risk-return relation is measured by π_P (π_N) under a prior positive (negative) return.

Table A.4
Estimation Results of the Intertemporal Risk-Return Relation under High/Low Investor Sentiment for the Equal-Weighted Market Portfolio

	Constant RRA under High/Low Market Sentiment		Prior 1-day Positive/Negative Price Changes		Prior 2-day Positive/Negative Price Changes		Prior 3-day Positive/Negative Price Changes		One Standard Dev. Positive/Negative Return Changes		Two Standard Dev. Positive/Negative Return Changes	
$c^H(\times 100)$	0.039	0.028	0.032	0.023	0.038	0.030	0.040	0.030	0.029	0.022	0.031	0.024
	(4.99)	(2.74)	(3.77)	(2.04)	(4.40)	(2.94)	(4.91)	(3.06)	(3.44)	(2.19)	(3.81)	(2.39)
$c^L(\times 100)$		0.051		0.042		0.046		0.050		0.037		0.039
		(4.34)		(3.24)		(3.63)		(4.10)		(2.58)		(2.89)
$\delta^H_{(P)}$	0.930	0.911	-1.424	-1.525	6.149	6.001	10.006	9.781	-2.068	-2.001	-3.079	-3.023
	(0.53)	(0.52)	(-0.51)	(-0.54)	(3.73)	(3.56)	(4.10)	(4.07)	(-0.53)	(-0.51)	(-0.62)	(-0.61)
δ^L_P			-3.447	-3.421	-3.060	-3.011	-0.084	-0.007	-2.908	-2.934	-3.114	-3.136
			(-3.10)	(-3.08)	(-1.07)	(-1.13)	(-0.04)	(0.00)	(-2.88)	(-2.90)	(-3.33)	(-3.35)
$\delta^H_{(N)}$	1.105	1.108	2.002	2.009	1.713	1.722	0.780	0.792	2.451	2.476	2.110	2.129
	(1.10)	(1.11)	(1.03)	(1.03)	(0.80)	(0.79)	(0.37)	(0.37)	(1.11)	(1.11)	(0.82)	(0.83)
δ^L_N			6.134	6.112	3.566	3.537	1.656	1.622	6.965	6.904	7.474	7.423
			(5.25)	(5.26)	(2.07)	(2.02)	(0.47)	(0.44)	(5.21)	(5.06)	(6.41)	(6.43)
$\phi(1)$	0.323	0.322	0.353	0.352	0.335	0.334	0.325	0.324	0.368	0.368	0.363	0.363
	(11.24)	(11.18)	(12.57)	(12.59)	(11.51)	(11.41)	(11.14)	(11.11)	(12.01)	(11.96)	(11.31)	(11.31)
$AdjR^2(\%)$	4.73	4.74	5.59	5.60	5.01	5.02	4.79	4.80	5.58	5.58	5.67	5.67

Note: This table reports estimates of the constant (Model 5) and asymmetric (Model 6) intertemporal risk-return relation under high/low sentiment for the equal-weighted market portfolio for the period over July 1965 – December 2018:

Model 5: Constant RRA under high/low sentiment

$$r_{m,t+1} = \left[c^H + \delta^H \hat{\sigma}^2_{m,t+1} + \phi_H(1) r_{m,t}\right] \cdot H + \left[c^L + \delta^L \hat{\sigma}^2_{m,t+1} + \phi_L(1) r_{m,t}\right] \cdot L + \varepsilon_{m,t+1},$$

Model 6: Asymmetric RRA under high/low sentiment

$$r_{m,t+1} = \left[c^H + (\delta^H_P Pd + \delta^H_N Nd)\hat{\sigma}^2_{m,t+1} + \phi_H(1) r_{m,t}\right] \cdot H + \left[c^L + (\delta^L_P Pd + \delta^L_N Nd)\hat{\sigma}^2_{m,t+1} + \phi_L(1) r_{m,t}\right] \cdot L + \varepsilon_{m,t+1},$$

where $r_{m,t+1}$ is daily realized excess return of the equal-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance. H (L) is the dummy representing high- (low-) sentiment regimes. Pd (Nd) is the dummy to capture not only prior d-day positive (negative) returns but also prior one and two standard deviation of positive and negative return changes. The RRA parameter in the high-sentiment regime is measured by δ^H_P (δ^H_N) under prior positive (negative) returns, while δ^L_P (δ^L_N) measures the RRA parameter under prior positive (negative) returns in the low-sentiment regime. The price adjustment during the high-sentiment regime is measured by $\phi_H(1)$, while it is measured by $\phi_L(1)$ during the low-sentiment regime.

Table A.5
The Asymmetric Intertemporal Risk-Return Relation Conditioned on Business Cycle for the Equal-Weighted Market Portfolio

	Constant RRA under Expansion/Recession in Business Cycle	Prior 1-day Positive/Negative Price Changes		Prior 2-day Positive/Negative Price Changes		Prior 3-day Positive/Negative Price Changes		One Standard Dev. Positive/Negative Return Changes		Two Standard Dev. Positive/Negative Return Changes	
$\mu^E_{(P)}(\times 100)$	0.021	0.022	0.055	0.023	0.048	0.035	0.059	0.035	0.053	0.043	0.078
	(2.49)	(2.55)	(4.47)	(2.61)	(2.90)	(4.81)	(3.59)	(5.47)	(1.09)	(6.75)	(0.57)
$\mu^R_P(\times 100)$			0.058		0.086		0.080		0.116		0.141
			(2.33)		(2.74)		(2.28)		(1.35)		(0.64)
$\mu^E_N(\times 100)$			-0.019		-0.049		-0.107		0.039		0.275
			(-1.08)		(-1.96)		(-2.76)		(0.75)		(2.60)
$\mu^R_{(N)}(\times 100)$	-0040	-0.041	-0.161	-0.035	-0.265	-0.011	-0.290	-0.024	-0.123	-0.009	-0.154
	(-1.69)	(-1.81)	(-4.68)	(-1.26)	(-4.94)	(-0.46)	(-3.59)	(-1.02)	(-1.36)	(-0.38)	(-0.96)
$\delta^E_{(P)}$	3.879	0.063	0.138	-0.116	0.785	-4.513	-3.788	-2.268	-2.309	-2.640	-3.451
	(3.80)	(0.04)	(0.08)	(-0.08)	(0.54)	(-1.74)	(-1.51)	(-1.14)	(-1.05)	(-0.82)	(-0.83)
δ^R_P		2.304	1.973	-0.132	-0.335	0.763	0.416	1.214	-0.007	1.819	0.274
		(1.74)	(1.35)	(-0.07)	(-0.19)	(0.29)	(0.15)	(0.73)	(0.00)	(1.15)	(0.14)
δ^E_N		7.005	6.882	10.877	11.041	14.755	17.005	8.715	8.423	8.512	6.561
		(3.91)	(3.66)	(3.82)	(3.52)	(4.55)	(4.21)	(4.57)	(4.22)	(3.85)	(3.11)
$\delta^R_{(N)}$	3.823	5.132	5.886	10.502	12.292	10.473	13.549	7.130	8.048	8.297	9.830
	(3.94)	(2.76)	(2.73)	(3.92)	(5.20)	(3.33)	(4.39)	(2.97)	(2.68)	(3.41)	(3.38)
$\phi(1)$	0.304	0.339	0.306	0.378	0.334	0.376	0.342	0.355	0.353	0.342	0.354
	(11.53)	(11.30)	(9.08)	(11.78)	(9.06)	(11.58)	(10.08)	(12.09)	(11.98)	(12.16)	(13.17)
$AdjR^2(\%)$	5.73	5.86	6.03	6.24	6.45	6.16	6.31	6.08	6.03	6.10	6.05

Note: This table reports the estimates of the constant (Model 7) and asymmetric (Model 8) intertemporal risk-return relation for the equal-weighted market portfolio conditioned on expansion and recession periods for the full period Jan. 2, 1926 – Dec. 31, 2019:

Model 7: Constant RRA under expansion/recession period

$$r_{m,t+1} = \left[\mu^E + \delta^E \hat{\sigma}^2_{m,t+1}\right] \cdot E + \left[\mu^R + \delta^R \hat{\sigma}^2_{m,t+1}\right] \cdot R + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1},$$

Model 8: Asymmetric RRA under expansion/recession in business cycle

$$r_{m,t+1} = \left[(\mu^E_P Pd + \mu^E_N Nd) + (\delta^E_P Pd + \delta^E_N Nd)\hat{\sigma}^2_{m,t+1}\right] \cdot E + \left[(\mu^R_P Pd + \mu^R_N Nd) + (\delta^R_P Pd + \delta^R_N Nd)\hat{\sigma}^2_{m,t+1}\right] \cdot R + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1},$$

where $r_{m,t+1}$ is daily realized excess return of the equal-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance. Pd (Nd) is the dummy to capture not only prior d-day positive (negative) returns but also prior one and two standard deviation positive and negative returns. E (R) is the dummy variable representing the expansion (recession) period in business cycle defined by the NBER. Periods of expansion begin at the trough date and end at the peak date, while periods of recession begin at the peak date and end at the trough date. The RRA parameter in the expansion periods is measured by δ^E_P (δ^E_N) under prior positive (negative) returns, while δ^R_P (δ^R_N) measures the RRA parameter under prior positive (negative) returns in the recession periods in business cycle.

Table A.6
Estimation Results of the Short-sale Effect of the Introduction of Options on Intertemporal Risk-Return Relation for the Equal-Weighted Market Portfolio

	1st Case of P and N: Prior 1-day positive-negative price change			2nd Case of P and N: One standard deviation return change			3rd Case of P and N: Two standard deviation return change		
	Model 9A	Model 9B	Model 9C	Model 9A	Model 9B	Model 9C	Model 9A	Model 9B	Model 9C
$c_{(1)}(\times 100)$	0.015	0.008	0.057	0.026	0.022	0.086	0.034	0.033	0.101
	(1.85)	(0.87)	(6.00)	(3.81)	(2.39)	(2.40)	(4.84)	(3.38)	(0.89)
$c_2\,(\times 100)$		0.017	-0.041		0.011	-0.001		0.002	0.120
		(1.11)	(-2.66)		(0.81)	(-0.02)		(0.18)	(1.34)
δ_P	2.132	2.263	2.344	1.407	1.458	1.032	2.808	2.817	1.921
	(1.56)	(1.64)	(1.68)	(0.88)	(0.92)	(0.64)	(1.60)	(1.61)	(1.05)
δ_N	5.189	5.300	5.205	7.380	7.432	7.429	7.441	7.450	6.928
	(3.87)	(3.93)	(3.74)	(4.49)	(4.49)	(4.29)	(4.70)	(4.69)	(4.10)
δ_P^O	-5.631	-6.115	-6.068	-9.150	-9.316	-9.441	-11.382	-11.407	-11.450
	(-3.75)	(-4.06)	(-3.87)	(-4.51)	(-4.61)	(-4.40)	(-4.75)	(-4.76)	(-4.53)
δ_N^O	2.474	2.105	2.713	1.134	0.980	1.209	2.678	2.655	2.583
	(0.82)	(0.67)	(0.84)	(0.30)	(0.25)	(0.31)	(0.60)	(0.59)	(0.60)
$\phi(1)$	0.341	0.340	0.310	0.354	0.354	0.350	0.338	0.338	0.350
	(10.56)	(10.58)	(10.06)	(11.70)	(11.72)	(11.26)	(12.58)	(12.57)	(13.23)
$AdjR^2(\%)$	5.90	5.90	6.03	6.20	6.20	6.17	6.29	6.29	6.22
$\delta_P+\delta_P^O$	-3.499	-3.852	-3.724	-7.743	-7.858	-8.409	-8.574	-8.590	-9.529
(t-value)	(-3.30)	(-3.47)	(-3.51)	(-6.62)	(-6.66)	(-6.04)	(-6.48)	(-6.48)	(-6.48)
$\delta_N+\delta_N^O$	7.663	7.405	7.918	8.513	8.412	8.638	10.119	10.105	9.511
(t-value)	(2.27)	(2.18)	(2.23)	(2.11)	(2.08)	(2.10)	(2.27)	(2.26)	(2.13)

Note: This table reports estimates that measure the effect of the introduction of options trading on the intertemporal risk-return relation. The following three models are estimated for the equal-weighted market portfolio for the full period Jan. 2, 1926 – Dec. 31, 2019:

Model 9A: $r_{m,t+1} = c + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta_P^O P + \delta_N^O N)\hat{\sigma}^2_{m,t+1} \cdot O + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 9B: $r_{m,t+1} = (c_1 + c_2 O) + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta_P^O P + \delta_N^O N)\hat{\sigma}^2_{m,t+1} \cdot O + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

Model 9C: $r_{m,t+1} = (c_1 P + c_2 N) + (\delta_P P + \delta_N N)\hat{\sigma}^2_{m,t+1} + (\delta_P^O P + \delta_N^O N)\hat{\sigma}^2_{m,t+1} \cdot O + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

where $r_{m,t+1}$ is daily realized excess return of the equal-weighted market portfolio, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance. O is the dummy variable to represent the period after the introduction of options into the markets (January 1981–December 2019). P (N) is the dummy to represent three different cases of state-dependent price dynamics. For the first case, P (N) captures prior 1-day positive (negative) return. For the second and third cases, it captures prior one and two standard deviation positive (negative) returns, respectively. δ_P^O (δ_N^O) measures the differential effect of the introduction of options on the intertemporal risk-return relation. $\delta_P+\delta_P^O$ ($\delta_N + \delta_N^O$) is the RRA parameter after the introduction of options. We also report the t-value for $H_0: \delta_P+\delta_P^O = 0$ and $H_0: \delta_N+\delta_N^O = 0$.

Table A.7
Estimation Results of the Intertemporal Risk-Return Relation Using the GARCH (1, 2) and TARCH (1, 2) Models

	Value-weighted Market Index Portfolio						Equal-weighted Market Index Portfolio					
	GARCH (1, 2) Model			TARCH (1, 2) Model			GARCH (1, 2) Model			TARCH (1, 2) Model		
	$d = 1$	$d = 2$	$d = 3$	$d = 1$	$d = 2$	$d = 3$	$d = 1$	$d = 2$	$d = 3$	$d = 1$	$d = 2$	$d = 3$
$c_P(\times 100)$	0.083	0.089	0.073	0.080	0.086	0.072	0.063	0.062	0.056	0.054	0.050	0.051
	(6.98)	(4.70)	(3.86)	(6.73)	(4.45)	(3.96)	(5.24)	(3.48)	(2.73)	(4.57)	(3.06)	(2.57)
$c_N(\times 100)$	-0.049	-0.047	-0.077	-0.049	-0.047	-0.079	-0.028	-0.059	-0.083	-0.026	-0.056	-0.091
	(-3.49)	(-2.12)	(-2.23)	(-3.59)	(-2.19)	(-2.32)	(-1.81)	(-2.55)	(-2.53)	(-1.68)	(-2.45)	(-2.92)
δ_P	-1.115	-4.899	-6.725	-0.926	-5.796	-8.305	1.268	0.471	-0.101	0.602	-0.276	-0.927
	(-1.07)	(-3.23)	(-3.21)	(-0.79)	(-3.71)	(-3.75)	(1.80)	(0.66)	(-0.09)	(0.70)	(-0.30)	(-0.69)
δ_N	2.405	5.718	11.077	2.328	5.652	10.235	3.309	7.434	10.145	4.351	8.177	10.349
	(1.73)	(2.77)	(3.29)	(1.72)	(2.83)	(3.29)	(2.65)	(3.76)	(3.91)	(3.58)	(4.69)	(5.46)
$\phi(1)$	0.043	0.094	0.117	0.048	0.103	0.126	0.274	0.311	0.323	0.310	0.347	0.344
	(1.67)	(2.94)	(3.66)	(1.80)	(3.12)	(3.81)	(8.12)	(7.79)	(8.06)	(8.93)	(9.07)	(8.90)
$AdjR^2(\%)$	1.01	1.19	1.35	1.01	1.29	1.50	5.77	6.27	6.09	5.91	6.51	6.29

Note: This table reports the estimates of Model 3 with the conditional variance estimated from the GARCH (1, 2) and the asymmetric Threshold-GARCH (1, 2) models for both the value-weighted and the equal-weighted market portfolios for the full period Jan. 2, 1926 – Dec. 31, 2019:

Model 3: $r_{m,t+1} = (c_P Pd + c_N Nd) + (\delta_P Pd + \delta_N Nd) \cdot \hat{\sigma}^2_{m,t+1} + \sum_{j=1}^{5} \phi_j r_{m,t+1-j} + \varepsilon_{m,t+1}$,

where $r_{m,t+1}$ is daily realized excess return of market portfolios, and $\hat{\sigma}^2_{m,t+1}$ is its conditional variance estimated from the GARCH (1, 2) and TARCH (1, 2) models. ϕ_j is the j^{th} order return autocorrelation coefficient, and $\phi(1)$ is the sum of autocorrelation coefficients. Pd (Nd) is the dummy to capture prior d-day positive (negative) returns, such that $P2 = 1$ when $e_{m,t-1} > 0$ and $e_{m,t} > 0$ ($N2 = 1$ when $e_{m,t-1} < 0$ and $e_{m,t} < 0$) where $e_{m,t}$ is the mean-deviated excess market returns. The RRA parameter is measured by δ_P (δ_N) under prior d-day positive (negative) returns. The numbers in parentheses are the Newey-West (1987) adjusted t-statistics. $Adj.R^2$ (%) is the percentage adjusted R^2.