The Term Structure of Short Selling Costs

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Abstract

I derive the term structure of short selling costs using the put-call parity relationship. The shape of the term structure is determined by informed investors' beliefs of when negative information will enter the market and correct the overpricing. I show that forward costs predicts future costs and stock returns, consistent with the expectations hypothesis in the model. I also find that an upward sloping curve around the earnings announcement increases the probability of a negative earnings surprise by 8.1%, supporting the prediction that short selling costs are higher when negative information is more likely to arrive.

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1 Introduction

If an investor believes an asset is overvalued, how does she determine how much she will pay to short it? I argue that time, not just overvaluation, is a critical component of equilibrium short selling costs. Will the next earnings announcement reveal information to the public that will correct the overvaluation? Perhaps the correction will take many months. As such, the market price for short selling should be directly related to when short sellers believe the correction will occur. This implies a dynamic relationship between shorting selling costs, expected returns and the timing of information in the market.

Miller (1977) presents a seminal theory in which stock prices can be elevated when there is disagreement among investors and pessimists are not able to trade based on their beliefs. However, in the equity markets, pessimists are generally able to short at an elevated cost rather than not short at all (Duffie, Garleanu and Pederson; 2002, Blocher, Reed and Van Wesep; 2013). It remains unclear how the price of the stock and the cost of shorting jointly change over time. An intrinsic problem is that if a stock is overpriced now, the price may rise before it falls (Scheinkman and Xiong; 2003); or, as in Diamond and Verecchia (1987), it may take a long time for the stock price to fully reflect private information. For an overpricing to be corrected, either short sale constraints must slacken, beliefs become less heterogenous, or both. New information can cause beliefs to converge and correct mispricings; however, information arrival is not necessarily a smooth process over time. Therefore, short selling costs will be higher when negative information is more likely to arrive. This relationship is difficult to distinguish empirically in one day maturity stock loan fees; however, a term structure of short selling costs can convey when pessimists believe the correction will occur.

To formalize my hypothesis, I present a model in which short sale constraints are formed when uninformed, optimistic investors have significant demand for an asset. These uninformed investors are agnostic or unaware of lending fees, and their demand only changes when new information arrives in the market.[1] Informed investors know the

[1]I do not distinguish between "soft" or "hard" information; however, Engelberg, Reed and Ringgenberg (2012) find that short sellers more often trade on the former.

true valuation but cannot fully influence the price because the uninformed restrict the supply of lendable shares. The valuation is revealed to all investors at a random date in the future, at which point the asset price drops to its valuation. Given the price of the asset, the beliefs of the informed investors determine the equilibrium short selling costs. Informed investors are able to short over multiple periods, which yields a simple equilibrium condition, or expectations hypothesis, in which the long-term rate of any stock loan is determined by expected future short-term lending fees. Similarly, the expected return each period equals the additive inverse of expected lending fees. The distribution of the information arrival process determines the shape of the term structure of shorting costs and expected returns.

To test the implications of the model, I exploit the fact that arbitrageurs can replicate a term short position by buying a put and shorting a call at the same strike and maturity. Using the put-call parity no-arbitrage condition, I derive term shorting costs and build a monthly, six month term structure by interpolating between available options. I find that forward option shorting costs are highly predictive of future option shorting costs as well as changes in option shorting costs. I also estimate Fama-Macbeth (1973) regressions, matching forward shorting costs to excess returns and find that option shorting costs predict negative cross-sectional excess returns over the six month horizon. The estimated coefficients of both the shorting cost and return predictability regressions are consistent with the expectations hypothesis in the model. In addition, forward shorting costs appear to have incremental return predictability for the corresponding month of returns beyond the first month shorting costs, providing evidence that forward costs are the more granular predictors of returns. In other words, when measuring return predictability of short selling costs, horizon matters.

Next, I relate the term structure of short selling costs to information arrival. Reed (2007) shows that short sale constrained stocks react more negatively to bad news, a phenomenon which suggests investors' beliefs can converge as information is revealed to the market. I thus test how the slope of the term structure around the earnings announcement relates to the probability of a negative earnings surprise. After controlling for the last option shorting cost prior to the earnings announcement, a positive difference between the shorting cost of the first option expiring after an earnings announcement and that of the last option expiring prior to the announcement (an upward sloping curve) increases the probability of a negative earnings surprise by 8.1% and leads to a -12bp CAR (-15.1% annually) over the earnings announcement period. Using the raw spread between the two options is also predictive of negative earnings announcements and abnormal returns. These findings suggest that short sellers pay more to short stocks over periods that include the earnings announcement when they believe there is a strong likelihood of a negative surprise. This result is consistent with the prediction of the model that shorting costs that vary over the term structure are due to time-varying information arrival probabilities. To my knowledge, this is the first paper to directly show that short selling costs are higher when negative information is more likely to arrive.[2]

Finally, I test the relationship between option shorting costs and the fees a short seller pays to borrow stock. Despite predicting levels, option shorting costs only minimally predict changes in stock loan fees, suggesting the expectations hypothesis does not hold across the options and stock loan market. However, option shorting costs have substantial incremental return predictability when both variables are included in Fama Macbeth (1973) regressions. Engelberg, Reed, and Ringgenberg (2016) find evidence that stock loan fees do not reflect the full cost of shorting due to volatility in the fee and the ability of lenders to recall loans at any time. This implies that the benefit of avoiding recalls may be capitalized into option prices. Thus, I regress the realization of a future recall, as defined as a 10% drop in lendable shares, on the difference between the option shorting cost and the current stock loan fee and find that the estimate is both statistically and economically significant after controlling for the current stock loan fee. This finding suggests that shorting in the options market may reflect the true, ex-ante cost of short selling over the term of the option.

[2] I do this by looking at shorting costs over different horizons within firm, calculated at a single point in time.

2 Description of the Stock Loan Market and Why the Term Structure Matters

For a more detailed overview of the securities lending market, I refer readers to Baklanova, Copeland and McCaughrin (2015). The main sources of lendable shares are brokers, institutional investors, such as mutual funds and pension funds, as well as ETF's. Hedge funds often borrow shares to execute short sales and form the main source of demand. To secure a stock loan, a borrower is required to post 102% cash collateral of the value of the underlying stock.[3] The borrower earns the federal funds rate on the cash collateral, but also pays the lender a fee to borrow the shares. The net interest rate is called the rebate. For stocks that are plentiful in supply to borrow, or general collateral, the lending fee is 25 basis points annualized or lower. For stocks that are more difficult to borrow and demand substantial, fees can exceed 10% annually.

Prime brokers are the main intermediaries matching supply of share lenders and demand of their hedge fund clients. Almost all stock loans are lent on an "open" basis where cash collateral and the stock loan fee are adjusted daily. Share lenders can terminate open stock loan arrangement at any time by issuing a recall. In particular, brokerages may be forced to recall shares if their retail clients sell shares or if there is an upcoming vote. Similarly, institutional investors may sell their shares currently being lent or recall them to exercise their ability to influence corporate actions (Aggarwal, Saffi and Sturgess; 2015).

Given that stock loan fees are regularly renegotiated, each day's fees likely reflect new economic or firm specific conditions. In the model, I will argue that the timing of the correction and the overvaluation jointly determine equilibrium lending fees. Thus it is difficult to disentangle these components empirically if the fee resets each day. Also, the information content of a stock loan fee is limited because it does not look forward beyond one day. For example, suppose a stock has an equilibrium one day lending fee of 1%. Assuming lending fees are determined as in the model described later, it would be impossible to distinguish whether the stock is 10% overvalued and has a 10% chance of

[3] All rules described are in the US. Retail investors need to post an additional 50% collateral.

being corrected or if it is 5% overvalued and has a 20% chance of being corrected on that day. Now suppose that short sellers can short for two periods and a stock has a lending fee of 1% for the first period and a 1.8% forward fee for the second period. Assuming lending fees are determined as in the model described in Section 4, the correction is more likely to occur in the second period. For instance, if the the stock is 10% overvalued the equilibrium fees would imply that the stock has a 10% chance of being corrected in the first period and a 20% chance in the second.[4] In reality, stocks probably do not experience corrections all at once; however, if at a particular point the term structure is upward sloping, it is reasonable to believe that negative news is more likely to come in the higher lending fee period.

3 Related Literature

The findings of this paper relate to several areas in the short selling, limits to arbitrage, bubbles and derivatives literature. Miller (1977) shows that investors who disagree and have downward sloping demand curves can cause overvaluation when short selling is disallowed. The theoretical portion of this paper most closely relates to two papers that jointly solve the price and equilibrium lending fee process: Blocher, Reed and Van Wesep (2013), henceforth BRVW, provide a simple framework where stock price and lending fee can simultaneously clear when long investors have downward sloping demand; and Duffie, Garleanu and Pederson (2002), henceforth DGP, present a dynamic model with disagreement among investors, search frictions, and an unknown day when all information is revealed.

Several papers have attempted to rationalize the seemingly irrational behavior of bubbles. Abreu and Brunnermeier (2003), Scheinkman and Xiong (2003) and Ofek and Richardson (2003) show that short sale constraints can lead to bubbles. Hong and Stein (2003) argue that bears' information is not initially revealed in market prices until the

[4] Period 1: $0.1 \cdot 0.1 = 1\%$, Period 2: $(1.0 - 0.1) \cdot 0.2 \cdot 0.1 = 1.8\%$. In the model, I assume the overvaluation is constant until it drops to 0. This may be unrealistic, but a more general identifying assumption for time-varying information arrival probabilities is that the expected overvaluation, conditional on the correction not occurring, is the current overvaluation. I show in the appendix that the equilibrium results and empirical predictions of the model hold with this more general assumption.

market begins a downturn. In Diamond and Verrecchia (1987) short sale constrained assets may be less informationally efficient and have more dramatic reactions to new information. Empirically, D'Avolio (2002) describes the market for borrowing stock and finds evidence that investor optimism limits arbitrage via the stock loan market.

Supporting Miller's hypothesis, several papers find evidence that short sale constraints are associated with overvaluation and lead to lower future returns (e.g., Jones and Lamont; 2002, Geczy, Musto and Reed; 2002, Nagel; 2005; and Drechsler and Drechsler; 2016). In contrast, Kaplan, Moskowitz and Sensoy (2013) find that exogenous increases in loan supply lead to reductions in lending fees, but have no impact on returns. Nagel (2005) proxies for short sale constraints with institutional ownership and finds that short sale constraints help explain cross-sectional asset pricing anomalies. This paper contributes to the literature on the relationship between short sale constraints and expected returns in that the term structure of option shorting costs provides a more dynamic measure of expected returns than the current stock loan fee or a single option maturity.

This paper also relates to short sale constraints spilling over to the options market. Ofek, Richardson and Whitelaw (2004) find that put-call parity deviations are higher among stocks that are expensive to borrow in the stock loan market, and the size of the deviation is a significant predictor of future negative returns.[5] Similarly, differences in implied volatility of calls and puts at the same strike and maturity are predictive of future returns (Cremers and Weinbaum; 2010). The option to stock volume ratio predicts future returns, suggesting that options are often used as an alternative to shorting in the options market (Johnson and So; 2012). Lending fees and option shorting costs are also tied to the value of a vote (Kalay, Karakas and Pant; 2014). Although this paper focuses on the asset pricing implications, the model does not view these explanations as mutually exclusive.[6]

Stock loan recalls can be a source of risk for short sellers and tend to occur on days

[5] A deviation being an implied stock price different than that predicted by put-call parity.

[6] The model is agnostic as to why uninformed demand is high and why uninformed do not lend all of their shares. It could be the case that uninformed investors value their votes which leads them to not lend all of their shares. This should still lead to the same asset pricing implications of the model because if it did not, investors could purchase stocks where the value to vote is high and lend them out to earn abnormal profits.

when the stock falls (D'Avolio; 2002). Engelberg, Reed and Ringgenberg (2016) show that this risk, along with volatility in loan market conditions are associated with lower expected returns. I find that the disparity in the price of shorting in the options market versus the stock loan market may be related to the risks of stock loans being recalled. These results are the first to directly compare option shorting costs and stock loan fees over the same horizon and are consistent with the findings of Engelberg, Reed and Ringgenberg (2016).

This paper also provides insight into the timing of information in financial markets. Swem (2016) finds that hedge funds anticipate information in markets in the form of analyst upgrades and downgrades. Among the findings specifically related to earnings announcements, short sale constrained stocks fall more after negative earnings surprises (Reed; 2007) and increases in short selling prior to earnings announcements are associated with informed traders anticipating a negative earnings surprise (Christophe, Ferri, and Angel; 2004). Prado, Saffi and Sturgess (2014) find that stocks with lower and more concentrated institutional ownership have smaller reactions on earnings days and greater post-earnings drifts. In other studies focusing on earnings announcements, Berkman, Dimitrov, Jain, Koch and Tice (2009) find that short sale constrained stocks with higher differences of opinion have more negative returns after earnings and Atilgan (2014) analyzes option prices around earnings and finds that implied volatility spreads between call and put options is predictive of returns around earnings. Compared to the other literature on option prices around earnings or information content in option prices, the empirical results do not rely on any specific options pricing model and assumptions underlying that model.

4 The Model

4.1 Model Overview

The model uses the equilibrium framework of BRVW to simultaneously solve price and shorting costs, or lending fees, in an infinite period setting.[7] The most important as-

[7]Short selling costs are equivalent to lending fees because there are no recalls or other frictions in the model.

sumption is that demand from long investors slopes downwards and they do not lend all of their shares in aggregate. I incorporate an information revealing process similar to to DGP, but I intentionally allow the arrival parameter to vary.[8] In DGP, investors search and bargain over stock loans. In this model, deep pocketed, informed investors competitively determine the equilibrium stock loan fees.[9] Unlike the aforementioned models, investors are able to short over multiple periods. There is no clear way of directly observing expected future lending fees from stock loan data; however, investors can replicate term stock loan arrangements in the options market. Because, the arrival parameter is allowed to vary, the shape of the lending fee curve can be informative as to when informed investors believe the correction will occur.

Several of the main equilibrium results could be reached using alternative frameworks and less restrictive assumptions; however, the intent of the model is to give clear and intuitive, testable predictions. These restrictions can be relaxed without materially altering the basic empirical predictions.

4.2 Model Basics

There is one asset that has a fixed number of shares N. The asset does not pay any dividends until a random, unknown time, τ, when the present value, V, of all future dividends is revealed to all investors. Prior to this day, no information is revealed regarding V and the price of the asset is P_t. There are two types of investors - informed and uninformed. Informed agents know V while uninformed agents make an inference on publicly available information of V. Informed investors agree on the probability distribution of τ and the independence of V and τ, while uninformed investors are entirely unaware of the information structure of the asset. Time is discrete and price and lending fees are determined in general equilibrium.

Short Selling: In each period, t, investors can short the asset for a length of time

[8] DGP acknowledge that the parameter could vary in their model but this is not their main focus.

[9] Search frictions and bargaining power would be difficult to measure empirically. The equilibrium is similar to the version of the BRVW model that is described in the appendix but there is an infinite number of periods and the information being revealed is random.

$k \in [1, T]$, with T being an arbitrarily large limit on the length of time investors can short.[10] In order to short, investors must borrow shares from investors who are long.[11] The fixed per-period, per-share, lending fee with time to maturity k, at period t, is $r_{t,k}$. When informed investors are long and lending fees are greater than 0, they lend all of their shares, $l^I = 1$; however, when uninformed investors are long, they always lend a fixed percentage of their shares $l^U \in [0, 1)$. In addition, uninformed investors do not change their demand of shares based on $r_{t,k}$.[12] Although this assumption may seem strange at first blush, there are institutional frictions that give evidence in its favor. For instance, brokerages are legally bound from lending the shares held by retail investors in cash accounts. Retail investors with margin accounts have their shares lent, but are not paid for the lending fees earned by the brokerage.[13] Similarly, investors at large institutions may have institution-wide lending programs where they are not aware of the fees they are earning or may be constrained in which stocks they can own (Evans, Ferreira and Prado; 2014). In fact, if every investor was willing and able to lend all of their shares short sale constraints would not exist.[14]

Investor Demand: Informed investors are risk-neutral and have unlimited access to capital. Let $R_t \equiv \{r_{t,1}, r_{t,2}, ...r_{t,T}\}$ denote the set of lending fees for all possible maturities. I define $D_t^I(P_t, R_t)$ as the aggregate number of shares informed investors demand for each price and set of lending fees pair. For a given R_t and P_t where expected profits for every possible shorting option is 0, the informed investor demand is flat. Decreasing any element of R_t causes informed investors to demand an infinite number of shares to short at that specific lending fee while increasing any element of R_t causes informed investors to demand an infinite number of shares to buy and lend at that lending fee.

[10] I disallow infinite period shorting because the cost of such shorting would be undefined and no such agreements exist in reality. I also disallow forward starting lending fee agreements purely for simplicity; however, forward lending fees can be inferred from simple lending fee contracts that begin at period t.

[11] There is no counterparty risk or discount rate in the model.

[12] BRVW show that this assumption is actually not needed in their more general model.

[13] A few brokerages have started "paid lending programs" where retail investors are paid for a portion of lending fees earned on their portfolio, but these are the exception rather than the norm.

[14] D'Avolio (2002) notes, "If all investors were institutionally and legally able to participate in lending, holding idle shares (i.e., not lending them) when fees are positive would be inconsistent with universal optimization and equilibrium."

Uninformed investors have limited capital and strictly downward sloping aggregate demand curves where $D_t^U(P_t) \geq 0$.[15] In equilibrium, the stock market must clear, therefore the shares demanded of the informed must equal the total shares outstanding minus the demand of the uninformed, $D_t^I = N - D_t^U$. If $D_t^U < N$ then $D_t^I > 0$ and informed investors are long so there is no short selling and the price equals the true valuation, $P_t = V$. When $D_t^U > N$, informed investors short sell the asset, $D_t^I < 0$, and available shares to borrow are $l^U D_t^U$.

Short Sale Constraint: An asset is short sale constrained if demand to short when lending fees are 0 is greater than the number of shares available to borrow, $-D_t^I(P_t, 0) > l^U D_t^U(P_t)$. Thus, lending fees must increase for the stock market to clear. If the asset is short sale constrained then it is overpriced, $P_t > V$, and all lending fees are strictly positive for $t < \tau$.[16] Figures 1 and 2 contain illustrations of how a short sale constraint arises. Once the information arrives in the market, the price equals the true valuation and all lending fees drop to 0.[17] For the rest of the paper I will focus on the scenario where there is a short sale constraint present prior to the revealing of the value of the asset.[18] The price of the asset does not change until information is revealed so:

$$P_{t+k} = \begin{cases} P_t & \text{if } t+k < \tau \\ V & \text{if } t+k \geq \tau \end{cases}$$

This is a stronger assumption than necessary. In the appendix I show that the equilibrium conditions are exactly the same when the expected price, conditional on no new information, is the current price. Uninformed demand could randomly change; however, it is simpler to proceed where the price only changes when information is revealed to the uninformed investors.

[15]For simplicity I assume that uninformed investors never short. Allowing them to short does not change the results of the model when there are short sale constraints. There would never be short sale constraints when the uninformed are short because market clearing requires $D^I > 0$ and $l^I = 1$.

[16]This also requires that the information arriving has above zero probability every period, $f_t(x = \tau | t < \tau) > 0 \ \forall t, x \geq t$, which I will assume later.

[17]Note that term lending fee contracts end at the specified maturity, not after the information arrives.

[18]If the asset is not short sale constrained there is no overpricing and $P_t = V$. To see more discussion of scenarios where the short sale constraint is slack, I refer readers to BRVW.

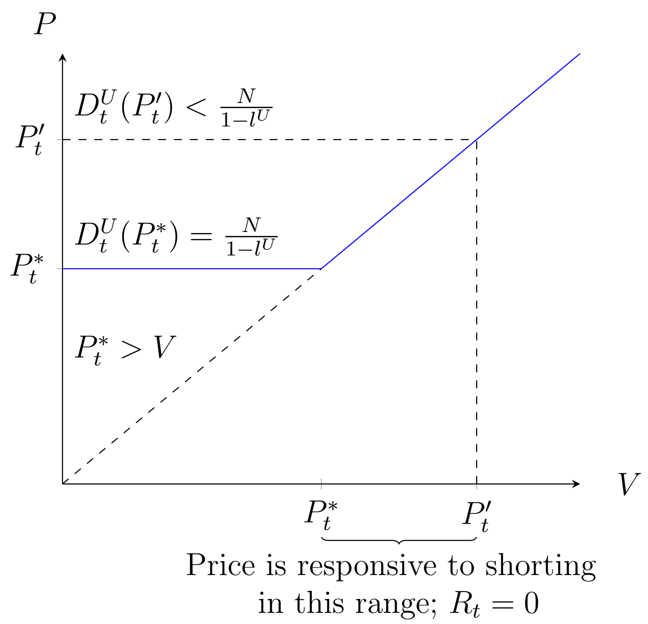

Figure 1. Relationship Between Price and Valuation. Price equals the true valuation when uninformed demand is less than total ownership $\frac{N}{1-l^U}$. In this range, the short sale constraint is slack and the price of the stock is sensitive to shorting and lending fees are 0. When uninformed demand equals total ownership, the price is greater than the true valuation if $-D_t^I(P_t, 0) > l^U D_t^U$. If this is the case, price is no longer affected by shorting and lending fees are strictly greater than 0.

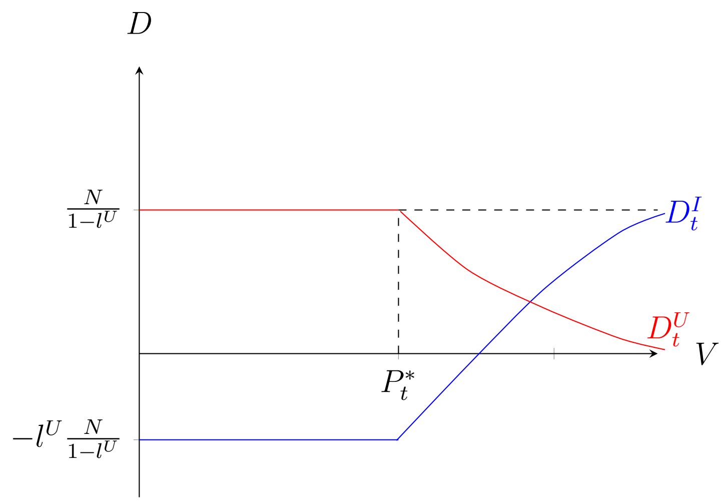

Figure 2. Investor Demand and Total Ownership. Shows the uninformed and informed demand at various valuations. When $V < P_t^*$, the short sale constraint is binding, uninformed investors demand the total ownership and informed investors are short all of the available shares to short. When V exceeds P_t^*, short selling decreases, total ownership decreases and uninformed demand decreases because of their downward sloping demand.

4.3 Equilibrium Lending Fees

All analysis hereafter on the distribution of τ, lending fees and any expectations are from the perspective of informed investors and assume that information has not yet arrived. Let λ_x denote the arrival parameter at time x and $\limsup_{x\to\infty} \lambda_x > 0$.[19] The cumulative density function of $\tau \le t+k$ at time t, is defined as:

$$F_t(t+k) = 1 - \prod_{x=t}^{t+k}(1-\lambda_x) \tag{1}$$

This is a very general arrival distribution that could take a number of forms.[20] I make no assumption of what are the "proper" $\lambda's$ are, but instead will use market prices to help infer their relative values empirically. The equilibrium condition of lending fees from period t to $t+k$ is the following:

$$\underbrace{r_{t,k}}_{\substack{\text{Per-period Lending}\\ \text{Fee}}} = \underbrace{\frac{r_{t,1} + \sum_{x=t+1}^{t+k} E_t[r_{x,1}]}{k}}_{\substack{\text{Average Expected}\\ \text{Lending Fee}}} = \underbrace{\frac{F_t(t+k)[P_t - V]}{k}}_{\substack{\text{Average Expected}\\ \text{Payoff}}} \tag{2}$$

The first two terms of the relationship are an expectations hypothesis: the rate informed investors pay to borrow stock over multiple periods should be equal to the average of the expected single period lending fees over the term of the agreement. Uninformed demand does not change until the information is revealed, thus $E_t[P_{t+k}] = VF_t(t+k) + (1 - F_t(t+k))P_t$, which gives the last term in the equilibrium condition. Note that in any competitive market with risk-neutrality, expected lending fees should equal the expected payoffs from shorting. Similarly, the per-period expected return of the asset also equals the additive inverse of the per-period lending fee. Thus stocks with higher lending fees have lower expected returns – a prediction that lends itself to empirical evidence.[21] This equilibrium condition establishes a clear relationship between differences in expected lending fees, expected returns and information arrival. Figure 3 displays an

[19]Mandelbaum, Hlynka and Brill (2007) show that the pdf of a nonhomogenous geometric distribution is nondefective when this condition holds. This assures that the information will always arrive.

[20]For instance in DGP, the model is solved with an exponentially distributed arrival process. Mandelbaum et al (2007) show that the nonhomogenous geometric distribution can describe all discrete distributions on non-negative integers.

[21]E.g. Jones and Lamont, 2002; Geczy, Musto and Reed, 2002 and Nagel, 2005.

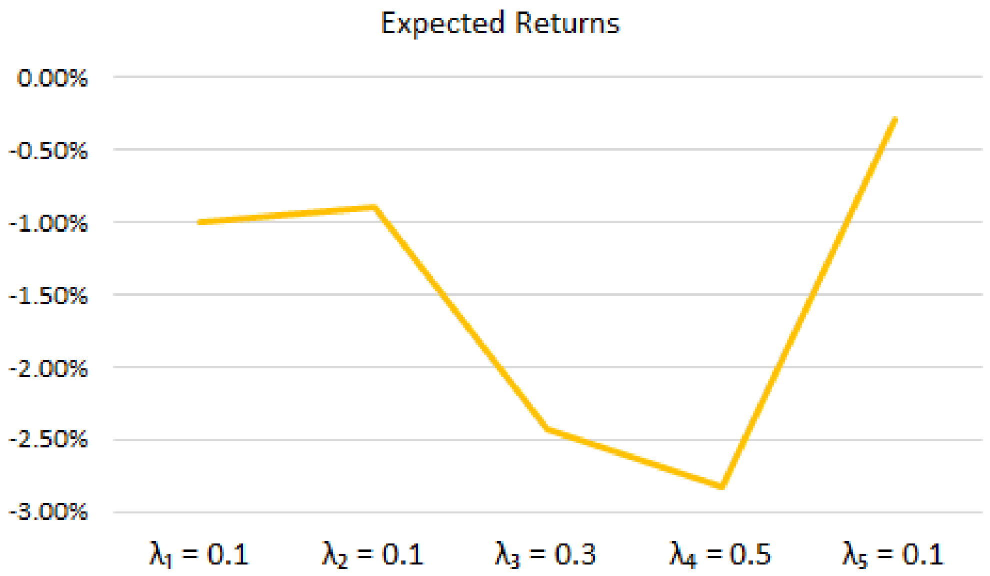

Figure 3. Term Structure of Expected Returns. Provides an illustration of the variation in per-period expected returns of a short sale constrained stock, at time t, over each period given a stock price normalized to \$10, a true valuation of \$9 and an arbitrary choice of λ_x for five periods. Expected returns equal the additive inverse of expected lending fees.

example of how expected returns vary over horizon based on the arrival parameter in each period. Each testable prediction presented in the next section immediately follow from equation (2).

4.4 Model Predictions

The following empirical predictions follow immediately from the main equilibrium condition and apply to overlapping or non-overlapping periods:

Prediction 1: Time intervals with higher per-period lending fees have higher average expected lending fees: if $r_{t,k} > r_{x,\gamma}$ then the expected per-period lending fee between t and $t+k$ is greater than the expected per-period lending fee between x and $x+\gamma$.

Prediction 2: Time intervals with higher per-period lending fees have lower average expected returns: if $r_{t,k} > r_{x,\gamma}$ then the expected per-period return between t and $t+k$ is lower than the expected per-period return between x and $x+\gamma$.

Prediction 3: Time intervals with higher per-period lending fees have a higher per-

period probability of information arriving: if $r_{t,k} > r_{x,\gamma}$, then the per-period probability of V being revealed between t and $t+k$ is higher than the per-period probability between x and $x+\gamma$. This empirical prediction applies within firm, not necessarily across firms as a firm with higher lending fees may just be more overvalued than another firm.

The first two predictions come directly from the expectations hypothesis, while the last prediction is related to the hypothesis that short selling costs are higher when information arrival is more likely. In the following sections I test each of these predictions empirically.

5 Empirical Analysis

I use data from Markit, OptionMetrics, IBES, CRSP and Compustat to test the predictions of the model.

5.1 Data

I obtain options data, dividend distributions and the zero coupon yield curve from OptionMetrics. As in Engelberg, Reed and Ringgenberg (2016), I drop option contracts with negative bid-ask spreads, bid-ask spreads that are greater than 50%, negative implied volatility, stock prices lower than \$5 or log moneyness greater than 0.5.[22] I also drop observations with dividend yields greater than 5%, annualized, over the life of the option. Following Ofek, Richardson and Whitelaw (2004), I calculate a term stock loan agreement with a fixed fee as described in the model as follows:[23]

$$S_{i,t}(1-v) = Call_{i,t} - Put_{i,t} + PV(K) + PV(D) \tag{3}$$

where *Call* is the mid-point of the closing call bid and ask prices for stock i at time t, *Put* the mid-point of closing put bid and ask prices, S the closing stock price, K the strike price of both the call and put, D the dividend paid over the life the option and T

[22]The results are robust to other filters.

[23]I do not incorporate the early exercise premium for calls as the fee that we are solving for is required to properly estimate it. By restricting the options to those closer to the money and with lower dividends the effect is likely mitigated. The results are very similar when all dividend payers are dropped.

the maturity date of both options. v represents the cost of shorting the over the term of the option, normalized by the stock price. I calculate v for every available option for each firm on each business day. There are generally multiple strikes available for each firm and maturity so I calculate the *Cost* as the median implied annualized cost across strikes for a specific maturity.

I match the options data at the firm-day level to equity lending data from Markit. The main variable of interest is *IndicativeFee*, which represents the estimation of the annualized interest rate a hedge fund would have to pay to borrow the stock on that day for one day. I also use the loan supply (*LendableQuantity*), which represents the total number of shares available to be lent. I match the data to CRSP, Compustat and IBES to obtain closing stock prices and earnings announcement data. As in Dellavigna and Pollet (2009), I use the earlier of the IBES and Compustat earnings announcement date and follow Johnson and So (2015) and drop earnings announcements if they are more than two days apart in the two databases. I restrict the sample to US equities. The resulting database includes data from July 1, 2006 to March 31, 2015.

To narrow the sample to short sale constrained stocks, I only include options with implied shorting costs greater than zero.[24] Negative shorting costs are likely caused by inaccurate or stale options prices and do not reflect short sale constrained stocks in which I am chiefly interested.[25]

To standardize the returns regressions, I transform the option shorting costs into one month maturity shorting costs. I require there be three or more option implied fees for each date firm pair so that there is a true term structure.[26] I linearly interpolate between all available options to obtain term shorting costs for every day before the end of six months. I then calculate one day forward costs using the interpolated costs and average the forward costs over the corresponding month of interest. In this manner, I create six one month maturity shorting costs that I define as $MonthCost_{i,t,m}$ where $m = 1 - 6$ and

[24]There are also some extreme positive values likely due to inaccurate option prices, so I trim the option implied fees at the 99.9% level.

[25]Negative shorting costs are theoretically unlikely as investors could short the forward and long the stock for a risk free profit.

[26]The results are not sensitive to this choice.

t is matched to the end of the month CRSP return dates. Every $MonthCost$ after month 1 is equivalent to a forward rate. For example, $MonthCost_{i,t,2}$ would be the forward shorting cost with a one month maturity that starts in one month.

Detailed summary statistics can be found in Panels A, B and C of Table 1 and the shorting costs are annualized. Panel D of the summary statistics includes the data that is used for the earnings announcement regressions, which I explain in further detail later. Notice that in Panel A, the mean and median option shorting cost slope downwards. This is consistent with arrival processes with persistence in the arrival parameter. For instance, the pdf form of equation (1) is decreasing in k if λ is fixed, which would lead to a downward sloping curve.[27] However, not all observations have downward sloping curves at all points in the term structure. In Section 5.3, I will show that observations with upward sloping curves around the earnings announcement have a higher likelihood of a negative earnings surprise.

5.2 Does the Option Shorting Cost Expectations Hypothesis Hold?

In this section, I first test whether option shorting costs predict future shorting costs (Prediction 1). I estimate a simple regression to test if period t forward costs predict period $t+m-1$, one month costs for $m = 2-6$ in the form of:

$$MonthCost_{i,t+m-1,1} = \beta_0 + \beta_1 MonthCost_{i,t,m} + \epsilon_{i,t,m} \tag{4}$$

If Prediction 1 holds, β_1 should be greater than 0 and a strict interpretation of the model would predict that $\beta_1 = 1$. All term structure regressions contain robust standard errors double clustered by date and firm following Petersen (2009). The estimated coefficients displayed in Table 2 are 0.87, 0.92, 1.00, 1.01 and 0.92 for $m = 2-6$, respectively and are all statistically significant.[28] It appears that forward option shorting costs predict

[27] $f_t(t+1) = \lambda > f_t(t+2) = (1-\lambda)\cdot\lambda > f_t(t+3) = (1-\lambda)^2\cdot\lambda$, etc.

[28] Throughout the paper, I refer to a result as statistically significant if it is statistically significant at the 10% level.

levels, which lends support to the expectations hypothesis. However, using differences is a more powerful test because levels are highly autocorrelated. I estimate the following term structure regression similar to Fama and Bliss (1987). The specification is the same as the previous, except I subtract the current one month shorting cost from both the dependent and independent variable for $m = 2 - 6$:

$$MonthCost_{i,t+m-1,1} - MonthCost_{i,t,1} = \beta_0 + \beta_1(MonthCost_{i,t,m} - MonthCost_{i,t,1}) + \epsilon_{i,t,m} \tag{5}$$

Similarly to the last regression, if Prediction 1 is true, β_1 should be greater than 0 and a strict interpretation of the model would predict that $\beta_1 = 1$. For $m = 2 - 6$, the estimated coefficients in Table 1 are 0.71, 0.79, 0.79, 0.80 and 0.80 and are all statistically significant. I include p-values to test the expectations hypothesis with a null hypothesis that $\beta_1 = 1$ and all are statistically different than 1. Because these are differences in costs measured using linear interpolation, there could be noise in the right hand side variables that could bias the coefficients downward. Nonetheless, given the sign and magnitudes of the coefficients, these results indicate that differences in current option shorting costs predict future changes in shorting costs and support the first empirical prediction of the model.

I now test if option shorting costs predict returns (Prediction 2). To do so, I use a regression approach similar to Boehmer, Jones, and Zhang (2008) and Engelberg, Reed and Ringgenberg (2016) to control for firm characteristics. More specifically, I estimate the following Fama-Macbeth (1973) regression for the beginning of each month in the sample for $m = 1 - 6$:

$$Ret_{i,t,m} = \beta_0 + \beta_1 MonthCost_{i,t,m} + \beta_2 Controls_{i,t,m} + \epsilon_{i,t,m} \tag{6}$$

where $Ret_{i,t,m}$ is the buy and hold return in excess of the one-month risk-free rate for the month following the current month. I include results with and without controls. The controls are the following: $Book/Market$ is the log of the book value of each firm divided by its market value, $MarketCap$, is the log of the firm's market capitalization,

$IdioVolatility$ is the log of idiosyncratic volatility calculated using the monthly standard deviation of the residual from a Fama-French three-factor regression, $Bid-Ask$ is the log of the closing bid-ask spread calculated as a fraction of the closing mid-point, and Ret_{t-1} is the stock return lagged by one month. The results are displayed in Table 3. The model predicts that β_1 should be negative and a strict interpretation of the model would predict that $\beta_1 = -1$. I find that for $m = 1 - 6$ with controls, the estimated coefficients are -1.05, -1.03, -1.00, -1.33, -1.53 and -1.41 and are statistically different than 0, while not statistically different from -1. These results are consistent with the model and provide evidence that forward shorting costs can predict excess returns several months in the future.

However, it could be the case that option shorting costs vary randomly across horizon and there is a common component that predicts returns over each month in the same way. To test if this is the case, I estimate the previous regression except I include the first month's option shorting cost, $MonthCost_{t,1}$, and the difference between the forward cost and first month cost, $MonthCost_{t,m} - MonthCost_{t,1}$, in every month specification for m = 2 - 6. According to the model, the forward cost should contain the relevant information for expected returns. Thus the model predicts that the coefficient for $MonthCost_{t,1}$ should be -1 as before, and the coefficient for $MonthCost_{t,m} - MonthCost_{t,1}$ should be negative (and -1 in a strict interpretation). The results are displayed in Table 4 and the estimated coefficients of $MonthCost_{t,1}$ with controls are all negative, statistically different from 0 while not statistically different than -1. All of the estimates of $MonthCost_{t,m} - MonthCost_{t,1}$ are negative with several statistically different from zero and none statistically different than -1. Given that forward costs seem to incrementally predict the corresponding month of returns beyond the first month option shorting cost, it appears that the term structure of shorting costs does explain the term structure of expected returns in the cross-section.

These initial tests of the predictiveness of changes in option shorting costs and excess returns lend support to an expectations hypothesis. Consistent with the model, option shorting costs can not only predict returns, but predict the term structure of stock returns.

Other factors that are related to the cross section of stock returns do not have this clear relationship with horizon.

5.3 Does the Term Structure Predict Negative Information Arriving in the Market?

I now test whether time-varying information arrival probabilities are related to the term structure of shorting costs (Prediction 3). Earnings announcements are potential event studies to test this prediction because new information is released to market participants. If short sellers are willing to pay more to short over a period that includes the earnings announcement date, then it is likely they believe there is a high probability that the earnings announcement will be negative and the stock price will fall. Also, higher shorting costs and lower expected returns around the earnings announcement are likely driven by the information being released in the earnings announcement, and not by expected changes in the stock price unrelated to this information.

For the tests in this section, I use daily frequency data, displayed in Panel D of Table 1, and the original option shorting costs rather than the constructed monthly ones. I also, cluster standard errors by firm and use earnings year-quarter fixed effects for all specifications.[29] It is not possible to infer the true option shorting costs on the day of the earnings announcement because options do not expire every day. However, it would still seem that if it is more expensive to short over a period that includes the earnings announcement than one that does not, the effect will still be present. Thus, I first estimate the following regression:

$$1_{[NegEarnings_i]} = \beta_0 + \beta_1 Cost_{i,t,\tau_i} + \beta_2 1_{[Cost_{i,t,T_i} - Cost_{i,t,\tau_i} > 0]} + \epsilon_{i,t} \tag{7}$$

where $1_{[NegEarnings_i]}$ is a dummy variable that equals 1 if the earnings is a negative surprise.[30] $Cost_{i,t,\tau_i}$ is the shorting cost of the last option expiring prior to the earnings

[29] There are multiple observations for the same earnings announcement at different points in time so the standard errors are likely to be biased downwards without clustering.

[30] More precisely, any SUE score, as defined by IBES, that is negative is treated as a negative earnings surprise.

announcement date and $1_{[Cost_{i,t,T_i} - Cost_{i,t,\tau_i} > 0]}$ is a dummy variable that equals 1 if the difference between the shorting cost of the first option expiring after the earnings announcement and that of the last option expiring before the announcement is greater than 0.[31] Figure 5 provides a visual description of the variables and Table 5 displays the results of this regression. In the sample, there is a 30.8% unconditional probability that a company reports a negative earnings surprise. A 1 percentage point increase in the annualized cost prior to the earnings announcement date, $Cost_{i,t,\tau_i}$, increases the probability of a negative earnings surprise by 0.62 percentage points (2.0%). More interestingly, a positive difference between shorting costs of the last option expiring before the earnings announcement date and the first option expiring afterwards increases the probability of a negative earnings surprise by 2.5 percentage points (8.1%) and is statistically significant. In terms of the model, these results can be interpreted as those companies with upwards sloping curves around the earnings announcements have relatively higher λ's for the earnings announcement day than days prior.

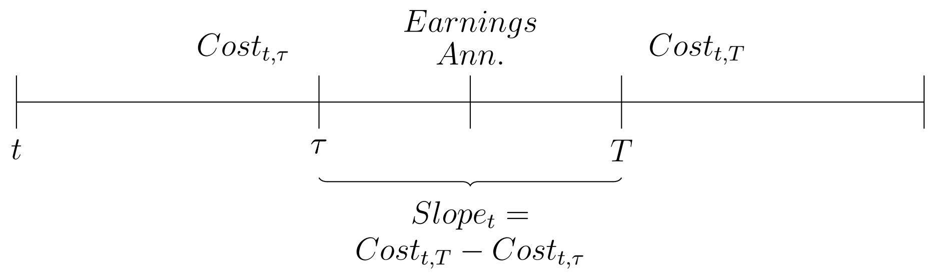

Figure 5. Timeline of Earnings Regressions. $Cost_{t,\tau}$ is the shorting cost of the last option to expire before the earnings announcement and $Cost_{t,T}$ is the shorting cost of the first option to expire after the earnings announcement. $Slope_t$ is the difference between $Cost_{t,\tau}$ and $Cost_{t,T}$.

In the next specification, also displayed in Table 5, I test if the magnitude of the slope is predictive of negative earnings surprises. I use the full sample, but also include specifications where I limit the sample to firms where the two options expire within 4, 2, and 1 week of each other, e.g. $T_i - \tau_i \leq 14$, and estimate the following regression:

$$1_{NegEarnings_i} = \beta_0 + \beta_1 Cost_{i,t,\tau_i} + \beta_2 Slope_{i,t} + \epsilon_{i,t} \tag{8}$$

[31] τ and T have subscripts for the i'th firm because they depend on the specific observation. The expiration dates of the options are not the same across firms.

In this regression $Slope_{i,t} = Cost_{i,t,T} - Cost_{i,t,\tau}$, which is the simple difference between shorting costs around the earnings announcement. This is an approximation for the slope of the term structure around the earnings announcement. Although this is not a perfect calculation, it should be more precise as the time between τ and T becomes smaller.[32] The estimated coefficient for $Cost_{i,t,\tau}$ is highly positive and significant and $Slope_{i,t}$ is also economically significant and statistically significant for all specifications. For instance, when limiting the sample to options that expire within a week of each other, a 1 percentage point increase in the option shorting cost increases the probability of a negative earnings surprise by 1.08 percentage points (3.8%), while a 1 percentage point increase in the slope, increases the probability of a negative earnings surprise by 0.77 percentage points (2.7%). Thus, it appears that the timing of information arrival is of first order importance for equilibrium short selling costs.

To ensure that slopes around earnings announcements are also associated with negative excess returns, I estimate the same regressions with the cumulative abnormal return during the window $[t, t+1]$ as the dependent variable. These results are displayed in Table 6. All of the abnormal returns are negative and a positive difference between the shorting costs leads to a statistically significant CAR of -12ps, which translates to -15.1% annually. These return tests provide additional support of Prediction 2 in that the slope of the term structure contains additional information regarding expected returns.

These results are agnostic to options pricing models and assumptions on the distribution of stock returns. Simple, no-arbitrage conditions predict negative earnings announcements. Market participants can observe options prices months in advance of the next earnings announcement and determine whether there is a high likelihood of a negative surprise.

[32] There is no way to know the exact shape of the lending fee curve between the two options.

5.4 The Relationship Between Option Shorting Costs and Stock Loan Fees

In this section, I analyze the relationship between option shorting costs and stock loan fees. For the remaining regressions, I use the constructed monthly term structure data. I first test whether option shorting cost levels can predict future daily stock loan fee levels and estimate the following regression for $m = 1 - 6$:

$$AverageFee_{i,t+m} = \beta_0 + \beta_1 MonthCost_{i,t,m} + \epsilon_{i,t,m} \tag{9}$$

where $AverageFee_{i,t+m}$ is the average, one month, realized daily stock loan fee, $IndicativeFee$, over month m. If option shorting costs and stock loan fees are completely fungible, the model would predict that β_1 should be $\beta_1 = 1$. The estimates of the coefficients, as displayed in Table 7, are 0.56, 0.85, 0.99, 1.16, 1.20 and 1.12 for $m = 1 - 6$, respectively and are statistically significant. The magnitude and significance of these coefficients suggest that option shorting costs and stock loan fees are closely related, which is consistent with the findings of Ofek, Richardson and Whitelaw (2004). However, it is not clear whether option shorting costs can predict changes in stock loan fees. I define $OptionPremium_{i,t,m} = MonthCost_{i,t,m} - IndicativeFee_{i,t}$; this represents the premium or discount of shorting in the options market versus shorting in the stock loan market. If these markets are completely fungible, then any premium or discount would reflect expected changes in stock loan fees over the corresponding month of the option shorting cost. To test if this is the case, I estimate the following regression:

$$AverageFee_{i,t+m} = \beta_0 + \beta_1 IndicativeFee_{i,t} + \beta_2 OptionPremium_{i,t,m} + \epsilon_{i,t,m} \tag{10}$$

If the option shorting cost predicts changes in realized stock loan fees, β_2 should be positive and the expectations hypothesis would predict a coefficient of 1. I find that the estimate of β_2 is statistically significant for all months, but not large (0.04, 0.11, 0.19, 0.27, 0.23, 0.18) as shown in Table 8. An explanation could be that $MonthCost$

is too noisy because it is calculated using options and the data comes from a different source. Therefore, I test if option shorting costs contain additional information regarding expected returns beyond the current stock loan fee. If this is the case, then noisy measurements of option shorting costs is unlikely to be a valid explanation for why option shorting costs do not seem to predict changes in stock loan fees.

I use the same specification as the previous returns regressions, but I include *IndicativeFee* and *OptionPremium* as the main independent variables. If option shorting costs have additional information regarding expected returns beyond the current stock loan fee, then the coefficient on *OptionPremium* should be negative. Indeed, as shown in Table 9, for all six months, with controls, the coefficients are negative (-0.77, -0.51, -0.52, -0.42, -1.47, -0.06), with months 1,2 and 5 statistically significant. These results indicate that option shorting costs contain additional information regarding expected returns beyond the current stock loan fee.

An alternative explanation for the apparent disconnect between shorting in these two markets could be that the stock loan market contains additional risks that the option market does not (Engelberg, Reed and Ringgenberg; 2016). Although the expectations hypothesis appears for the most part to hold within the options market, it may not across the options and stock loan market. A feature of the stock loan market is that loans can be recalled at any time by a lender; thus, stock recalls can be a source of risk for an arbitrageur as she may lose profitable shorting opportunities or if she is forced to cover her short position at a temporarily higher price (D'Avolio; 2002). In contrast, recalls are not possible in the options market because options are term derivative agreements.

Engelberg, Reed and Ringgenberg (2016) find that stocks that are more likely to be recalled have lower expected returns. I thus test whether recalls can be predicted by the relative price of shorting in the options market versus the stock loan market. More specifically, I test if the difference between the option shorting cost and stock loan fee predicts recalls beyond the current stock loan fee. I use a similar specification to Engelberg, Reed and Ringgenberg (2014) and define a stock recall as a 10% or greater reduction in lendable shares, *LendableQuantity*, in a month. In the sample, the probability of a recall over the next month is 4.2%. I estimate the following regression for $m = 1 - 6$:

$$1_{Recall_{i,t+m}} = \beta_0 + \beta_1 IndicativeFee_{i,t} + \beta_2 OptionPremium_{i,t,m} + \beta_3 Controls_{i,t} + \epsilon_{i,t,m} \quad (11)$$

where $1_{Recall_{i,t+m}}$ is an indicator variable that equals 1 if a recall occurs in month m. As before, $OptionPremium_{i,t,m} = MonthCost_{i,t,m} - IndicativeFee_{i,t}$. Among the controls, $MarketCap_i$ is the log of market capitalization, $Ret_{i,t-1}$ is the return lagged by one month and $Volume_{i,t-1}$ is the total trading volume divided by shares outstanding lagged by one month. I include date fixed effects and cluster by date to control for unobserved market-wide or macroeconomic shocks that could systematically affect recalls. The results are displayed in Table 10.

The coefficient for $IndicativeFee$ is statistically significant and large for all specifications as expected. When including controls, $OptionPremium$ is positive and statistically significant for all months. The economic magnitude is large as well. For instance, in the month one specification with controls, a 1 percentage point increase in the current lending fee increases the probability of a recall by 4.2%, while a 1 percentage point increase in the difference between the option shorting cost and current lending fee increases the probability of a recall by 1.9%. This suggests arbitrageurs actively pay premiums to avoid the risk of recalls and offers a potential explanation as to why option shorting costs have minimal incremental predictive power for stock loan fees. The model I present does not allow for recalls, but one can imagine if share lenders are given an option to recall the loan at any time, that the cost of this option will be reflected in lower equilibrium stock loan fees.

It is worth noting that the cause of lower equilibrium lending fees is not necessarily risk aversion. I do not claim that short sellers are perfectly risk-neutral, but based on the evidence in this paper, it does seem that risk-neutrality is a reasonable assumption for the relative price of shorting across horizon in the options market. That stock loan fees do not reflect the full cost of shorting may be a risk-aversion story driven by volatility in fees or a friction that reduces the expected value of the short position such as recalls, or a combination of both.

Consistent with the hypothesis of Engelberg, Reed and Ringgenberg (2016), it appears that the fees paid in the stock loan market do not reflect the full costs of short selling. My findings suggest that option shorting costs may reflect the true ex-ante cost of shorting over the maturity of the option. However, based on the earlier tests in this paper, it appears that horizon and information arrival probabilities are important components of equilibrium short selling costs within the options market.

6 Conclusion

Short sellers are concerned with overvaluation, but also when the overvaluation will be corrected; thus, short selling costs will be higher over periods when negative information is more likely to arrive. With this intuition, I build a model of information arrival that endogenizes price and shorting costs using a simple, zero profits equilibrium condition. I empirically test the predictions of the model by constructing term short agreements using the put-call parity no-arbitrage condition. I find that option shorting costs predict future costs and returns, results consistent with the expectations hypothesis of the model. I also use earnings announcements as an event study and find that upward sloping curves around earnings announcements predict negative earnings surprises as well as negative excess returns. Thus it appears that horizon is a critical component of equilibrium shorting costs. I also find that option shorting costs only minimally predict changes in stock loan fees despite containing additional information regarding expected returns, a rejection of the expectations hypothesis across markets. As a potential explanation, I find that the difference short sellers pay in the option markets relative to the stock loan market is predictive of a proxy for stock loans being recalled. These results are among the first in the empirical literature to capture the dynamic effects of short sale constraints and provide a new methodology for measuring shorting costs across different horizons. Short selling is rarely completely restricted in financial markets; therefore understanding how short sale constraints are priced across horizon is critical to understanding when information enters asset prices and how long mispricings persist.

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Table 1
Summary Statistics

Dates are matched to CRSP end of month return dates and options data comes from OptionMetrics. Panel A includes all one month option shorting costs greater than 0 with more than 3 options that pass the initial filters for months 1 - 6. Months 2 - 6 are forward rates. Panel B includes the average realized daily stock loan fee from Markit for the corresponding months. Panel C includes a correlation matrix for the given variables where L is an abbreviation for LoanFee and M is an abbreviation for MonthCost. All percentages are annualized and subscripts reference the month number. MonthCost is calculated at time t while LoanFee represents realizations in month m.

Panel A: Option Shorting Cost Statistics

Variable	Mean	Median	1st	99th	Standard Deviation	N
$MonthCost_1$	349.4 bps	164.4 bps	2.7 bps	2958.2 bps	631.7 bps	63071
$MonthCost_2$	237.6 bps	106.6 bps	2.6 bps	2272.8 bps	455.5 bps	55161
$MonthCost_3$	193.5 bps	91.5 bps	2.2 bps	1853.3 bps	370.4 bps	55135
$MonthCost_4$	169.0 bps	79.3 bps	1.7 bps	1590.8 bps	311.2 bps	51009
$MonthCost_5$	158.9 bps	77.5 bps	1.5 bps	1394.7 bps	272.4 bps	48480
$MonthCost_6$	159.2 bps	80.9 bps	1.6 bps	1293.0 bps	256.1 bps	48186

Panel B: Stock Loan Realized Fee Statistics

Variable	Mean	Median	1st	99th	Standard Deviation	N
$LoanFee_1$	163.1 bps	39.8 bps	25.6 bps	2913.0 bps	577.5 bps	60120
$LoanFee_2$	173.8 bps	39.8 bps	25.6 bps	3039.1 bps	614.7 bps	52131
$LoanFee_3$	172.3 bps	39.7 bps	25.6 bps	3023.8 bps	617.9 bps	51731
$LoanFee_4$	176.8 bps	39.8 bps	25.6 bps	3087.0 bps	635.8 bps	47464
$LoanFee_5$	179.2 bps	39.8 bps	25.6 bps	3090.9 bps	647.8 bps	44709
$LoanFee_6$	178.1 bps	39.7 bps	25.6 bps	3095.5 bps	644.4 bps	43683

Panel C: Correlation Matrix

	M_1	M_2	M_3	M_4	M_5	M_6	L_1	L_2	L_3	L_4	L_5	L_6
M1	1.000											
M2	0.852	1.000										
M3	0.837	0.894	1.000									
M4	0.759	0.826	0.921	1.000								
M5	0.634	0.684	0.776	0.850	1.000							
M6	0.673	0.736	0.763	0.817	0.862	1.000						
L1	0.763	0.775	0.767	0.733	0.633	0.646	1.000					
L2	0.717	0.721	0.719	0.689	0.593	0.607	0.944	1.000				
L3	0.653	0.661	0.668	0.645	0.553	0.568	0.868	0.946	1.000			
L4	0.601	0.606	0.618	0.602	0.521	0.537	0.798	0.867	0.944	1.000		
L5	0.547	0.562	0.570	0.557	0.488	0.503	0.744	0.802	0.872	0.950	1.000	
L6	0.506	0.522	0.527	0.517	0.454	0.470	0.700	0.748	0.809	0.880	0.953	1.000

Table 1 Cont'd
Summary Statistics

Panel D includes daily option shorting costs greater than 0. $Cost_{\tau_i}$ is the last option expiring before the next earnings announcement and $Cost_{T_i}$ is the first option expiring after the next earnings announcement. $Slope = Cost_{T_i} - Cost_{\tau_i}$. $PositiveSlope$ is a dummy variable that equals 1 if $Slope$ is positive. $NegEarnings$ is a dummy variable that equals 1 if the IBES earnings announcement comes in below analyst estimates (a negative SUE score). $Maturity$ is the number of days until the expiration of the corresponding option.

Panel D: Earnings Option Shorting Cost Statistics

Variable	**Mean**	**Median**	**1st**	**99th**	**Standard Deviation**	**N**
Cost_{τ_i}	380.0 bps	233.4 bps	5.4 bps	2472.2 bps	498.5 bps	1287971
Cost_{T_i}	270.4 bps	183.8 bps	6.9 bps	1675.4 bps	324.5 bps	
Slope	-138.0 bps	-57.4 bps	-1509.1 bps	312.2 bps	412.8 bps	
PositiveSlope	32.1%	0.0%	0.0%	100.0%	46.7%	
Negative Earnings	30.8%	0.0%	0.0%	100.0%	46.2%	
Days to Earnings	63.6	62.0	14.0	165.0	29.9	
Maturity_{τ_i}	44.9	43.0	8.0	128.0	23.6	
Maturity_{T_i}	107.0	108.0	36.0	219.0	46.9	

Table 2
Predictability of Future Shorting Costs

This table contains results testing whether forward option shorting costs predict future option shorting costs. I suppress all i subscripts indicating firm for ease of notation. $Y_{t,m}$ is the period t, one month, forward shorting cost corresponding to month m. $Y_{t+m-1,1}$ is the one month shorting cost starting at period $t+m-1$. The top regressions test if the current forward rate predicts the future one month rate for $m = 2-6$. The bottom regressions are testing differences by subtracting $Y_{t,1}$ from the left and right hand side variables. A strict interpretation of the model would predict a coefficient of 1 for b in both regressions. Robust standard errors double clustered by date and firm are shown below the parameter estimates in parenthesis. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively. I present p-values for the expectation hypothesis where $b = 1$.

	$Y_{t+m-1,1} = a + bY_{t,m} + e$				
m	2	3	4	5	6
$\hat{b}$	0.868***	0.916***	0.998***	1.013***	0.924***
	(0.030)	(0.048)	(0.058)	(0.071)	(0.075)
$\hat{a}$	0.010***	0.013***	0.014***	0.014***	0.016***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Exp. hyp. p-value	0.000	0.081	0.979	0.855	0.307
N	46175	44912	41061	37620	36163
R^2	0.375	0.283	0.225	0.174	0.135

	$Y_{t+m-1,1} - Y_{t,1} = a + b(Y_{t,m} - Y_{t,1}) + e$				
m	2	3	4	5	6
$\hat{b}$	0.706***	0.785***	0.789***	0.799***	0.800***
	(0.037)	(0.031)	(0.024)	(0.029)	(0.032)
$\hat{a}$	0.005***	0.008***	0.010***	0.011***	0.011***
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)
Exp. hyp. p-value	0.000	0.000	0.000	0.000	0.000
N	46170	44888	41041	37600	36136
R^2	0.189	0.276	0.291	0.307	0.327

Table 3
Cross-Sectional Return Predictability of Shorting Costs

This table contains Fama-Macbeth (1973) results testing the relationship between option shorting costs and returns. I suppress all i subscripts indicating firm for ease of notation. Ret_{t+m} is the buy and hold return in excess of the one-month risk-free rate in month m, $MonthCost_{t,m}$ is the option shorting cost for month m, $Book/Market$ is the log of the book/market ratio from Compustat, $MarketCap$ is the log of the market capitalization, $Idio.\ Volatility$ is the log of idiosyncratic volatility calculated using the monthly standard deviation of the residual from a Fama-French three-factor regression, $Bid - Ask$ is the log of the bid ask ratio and $Return_{t-1}$ is the return lagged by one month. A strict interpretation of the model would predict that the coefficient for $MonthCost$ would be -1. I report the time-series mean of the parameter estimates with t-statistics, calculated using Newey-West (1987) standard errors with three lags, shown below in parenthesis. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively. I present p-values for the expectation hypothesis where the coefficient of $MonthCost_{t,m}$ equals -1.

Explanatory Variable	**Monthly Excess Return$_{t+m}$**											
Month	**1**	**1**	**2**	**2**	**3**	**3**	**4**	**4**	**5**	**5**	**6**	**6**
MonthCost$_{t,m}$	-1.075***	-1.049***	-1.254***	-1.032***	-1.587***	-1.002***	-1.607***	-1.328**	-2.451***	-1.528***	-1.718***	-1.407**
	(-5.74)	(-4.72)	(-4.16)	(-3.55)	(-4.04)	(-2.82)	(-2.72)	(-2.47)	(-3.86)	(-3.33)	(-2.92)	(-2.42)
Book/Market		0.000		0.000		0.000		-0.000		-0.001		-0.001
		(0.01)		(0.16)		(0.00)		(-0.28)		(-0.72)		(-0.62)
MarketCap		0.001		0.001		0.001		0.001		-0.000		0.001
		(0.89)		(0.67)		(1.09)		(1.24)		(-0.26)		(0.59)
Idio. Volatility		-0.003		-0.007		0.002		-0.009		-0.018**		-0.003
		(-0.43)		(-1.09)		(0.30)		(-1.16)		(-2.29)		(-0.42)
Bid-Ask		0.000		0.000		-0.001		0.001		-0.002		0.001
		(0.13)		(0.20)		(-0.94)		(0.91)		(-1.44)		(0.86)
Return$_t$		-0.001		-0.004		-0.009		0.004		0.014		0.019*
		(-0.07)		(-0.36)		(-0.91)		(0.52)		(1.18)		(1.81)
Return$_{t-1}$		-0.013		-0.002		0.008		0.019		-0.000		0.012
		(-1.17)		(-0.19)		(1.41)		(1.60)		(-0.03)		(1.27)
Constant	0.008	-0.007	0.008	-0.004	0.009	-0.024	0.007	-0.011	0.006	-0.001	0.004	-0.006
	(1.21)	(-0.31)	(1.21)	(-0.16)	(1.33)	(-1.03)	(1.11)	(-0.47)	(1.01)	(-0.06)	(0.64)	(-0.24)
Exp. hyp. p-value	0.690	0.827	0.401	0.913	0.139	0.995	0.307	0.543	0.024	0.252	0.225	0.486
N	62955	49267	54894	42878	54660	42877	50310	39445	47542	37416	47002	37172
Average R^2	0.010	0.072	0.013	0.077	0.011	0.075	0.013	0.082	0.014	0.077	0.010	0.080

Table 4
Cross-Sectional Return Predictability of Shorting Costs

This table contains Fama-Macbeth (1973) results testing if the month m option shorting cost has return predictability beyond the month 1 cost. I suppress all i subscripts indicating firm for ease of notation. Ret_{t+m} is the buy and hold return in excess of the one-month risk-free rate in month m, $MonthCost_{t,m}$ is the option shorting cost for month m, $Book/Market$ is the log of the book/market ratio from Compustat, $MarketCap$ is the log of the market capitalization, $Idio.\ Volatility$ is the log of idiosyncratic volatility calculated using the monthly standard deviation of the residual from a Fama-French three-factor regression, $Bid-Ask$ is the log of the bid ask ratio and $Return_{t-1}$ is the return lagged by one month. I report the time-series mean of the parameter estimates with t-statistics, calculated using Newey-West (1987) standard errors with three lags, shown below in parenthesis. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively. I present p-values for tests where the coefficients of $MonthCost_{t,1}$ and $MonthCost_{t,m} - MonthCost_{t,1}$ equal -1.

Explanatory Variable	**Monthly Excess Return$_{t+m}$**									
Month	**2**	**2**	**3**	**3**	**4**	**4**	**5**	**5**	**6**	**6**
$MonthCost_{t,1}$	-1.260***	-1.120***	-1.458***	-1.038***	-1.642***	-1.489***	-2.355***	-1.609***	-1.595***	-1.345**
	(-3.98)	(-3.53)	(-3.53)	(-2.71)	(-2.74)	(-2.74)	(-4.03)	(-3.71)	(-2.75)	(-2.16)
$MonthCost_{t,m}$ - $MonthCost_{t,1}$	-0.637	-0.268	-1.109***	-0.559	-1.560**	-1.635**	-2.220***	-1.867***	-1.486**	-1.374**
	(-1.39)	(-0.57)	(-3.17)	(-1.48)	(-2.38)	(-2.46)	(-3.68)	(-3.24)	(-2.42)	(-2.03)
Book/Market		0.000		-0.000		-0.001		-0.001		-0.001
		(0.05)		(-0.07)		(-0.42)		(-0.65)		(-0.50)
MarketCap		0.000		0.001		0.001		-0.000		0.001
		(0.49)		(0.99)		(1.09)		(-0.25)		(0.62)
Idio. Volatility		-0.006		0.002		-0.009		-0.020**		-0.003
		(-0.91)		(0.30)		(-1.19)		(-2.44)		(-0.42)
Bid-Ask		0.001		-0.001		0.001		-0.002		0.001
		(0.53)		(-0.85)		(0.88)		(-1.26)		(0.79)
$Return_t$		-0.004		-0.009		0.006		0.013		0.019*
		(-0.30)		(-0.91)		(0.68)		(1.06)		(1.88)
$Return_{t-1}$		-0.002		0.008		0.020		0.000		0.012
		(-0.24)		(1.45)		(1.63)		(0.05)		(1.24)
Constant	0.008	0.002	0.009	-0.020	0.007	-0.009	0.007	0.000	0.004	-0.007
	(1.26)	(0.09)	(1.37)	(-0.87)	(1.10)	(-0.38)	(1.03)	(0.02)	(0.66)	(-0.28)
coeff($MonthCost_{t,1}$) = -1	0.413	0.707	0.270	0.922	0.287	0.370	0.022	0.163	0.308	0.581
coeff($MonthCost_{t,m}$ - $MonthCost_{t,1}$) = -1	0.432	0.122	0.757	0.246	0.395	0.341	0.046	0.136	0.431	0.582
N	54886	42872	54629	42852	50282	39423	47512	37393	46961	37142
Average R^2	0.018	0.084	0.017	0.079	0.020	0.088	0.021	0.086	0.020	0.090

Table 5
The Relationship Between the Slope of the Shorting Cost Curve around the Earnings Announcement and Negative Earnings Surprises

This table contains results testing whether differences in option shorting costs around the earnings announcement predict negative earnings surprises. I suppress all i subscripts indicating firm for ease of notation. Earnings announcement data comes from IBES and daily frequency data is used. $NegEarnings$ is a dummy variable that equals 1 if the earnings comes in below consensus analyst expectations (a negative SUE score). $Cost_{\tau}$ is the last option expiring before the next earnings announcement and $Cost_{T_i}$ is the first option expiring after the next earnings announcement. $Slope = Cost_T - Cost_{\tau}$. $PositiveSlope$ is a dummy variable that equals 1 if $Slope$ is positive. $Slope$ is included in regressions for observations where the two options expire within 4, 2, and 1 week of each other and without any restrictions. Earnings year-quarter fixed effects are used throughout and t-statistics are shown below the parameter estimates in parenthesis and are calculated using robust standard errors clustered by firm. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

Explanatory	**NegEarnings**				
Variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**
$Cost_{\tau}$	0.620***	0.999***	0.989***	1.085***	1.081***
	(11.80)	(10.57)	(9.08)	(3.20)	(3.11)
PositiveSpread	0.025***				
	(7.09)				
Spread		0.888***	0.825***	0.811**	0.766*
		(7.90)	(6.63)	(2.09)	(1.81)
Days Between Options			<= 28	<= 14	<= 7
Mean Dep. Var	0.308	0.308	0.308	0.294	0.288
N	1287971	1287971	312379	10540	8042
R^2	0.011	0.012	0.013	0.019	0.022

Table 6
The Relationship Between the Slope of the Shorting Cost Curve around the Earnings Announcement and Abnormal Returns

This table contains results testing whether differences in option shorting costs around earnings announcements predicts negative excess returns. I suppress all i subscripts indicating firm for ease of notation. Earnings announcement data comes from IBES and daily frequency data is used. $Cost_{\tau}$ is the last option expiring before the next earnings announcement and $Cost_T$ is the first option expiring after the next earnings announcement. $Slope = Cost_T - Cost_{\tau}$. $PositiveSlope$ is a dummy variable that equals 1 if $Slope$ is positive. CAR is the sum of the daily returns of the stock minus the market return over the period $[t, t+1]$ where t is the earnings announcement date. $Slope$ is included in regressions for observations where the two options expire within 4, 2, and 1 week of each other and without any restrictions. Earnings year-quarter fixed effects are used throughout and t-statistics are shown below the parameter estimates in parenthesis and are calculated using robust standard errors clustered by firm. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

Explanatory	**CAR[0,1]**				
Variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**
Cost_{τ}	-0.037***	-0.061***	-0.066***	-0.119***	-0.149***
	(-3.75)	(-3.17)	(-3.17)	(-2.64)	(-3.13)
PositiveSlope	-0.0012**				
	(-2.40)				
Slope		-0.053**	-0.061***	-0.047	-0.067
		(-2.39)	(-2.59)	(-0.88)	(-1.27)
Days Between Options			<= 28	<= 14	<= 7
N	1287763	1287763	312352	10540	8042
R^2	0.003	0.003	0.005	0.023	0.033

Table 7
Relationship Between Option Shorting Costs and Stock Loan Fees

This table contains results testing if option shorting costs predict realized stock loan fees. I suppress all i subscripts indicating firm for ease of notation. $IndicativeFee$ is the one day stock loan fee from Markit. $Realized\ Average\ Loan\ Fee_{t+m}$ is the one month realized average $IndicativeFee$ in month m. $MonthCost_{t,m}$ is the one month, option shorting cost at time t in month m. Months $m = 1 - 6$ are shown. Robust standard errors double clustered by date and firm are shown below the parameter estimates in parenthesis. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively. I present p-values for the expectation hypothesis where $MonthCost_{t,m} = 1$.

Explanatory Variable	**Realized Average Loan Fee$_{t+m}$**					
Month	**1**	**2**	**3**	**4**	**5**	**6**
MonthCost$_{t,m}$	0.557***	0.851***	0.989***	1.160***	1.197***	1.124***
	(0.056)	(0.072)	(0.082)	(0.092)	(0.103)	(0.107)
Constant	-0.003**	-0.003**	-0.002*	-0.002*	-0.001	0.000
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Exp. hyp. p-value	0.000	0.038	0.895	0.082	0.055	0.249
N	60120	52131	51731	47464	44709	43683
R^2	0.360	0.389	0.345	0.312	0.242	0.191

Table 8

Relationship Between Option Shorting Costs and Stock Loan Fees

This table contains results testing if option shorting costs provide incremental predictiveness of realized stock loan fees beyond the current stock loan fee. I suppress all i subscripts indicating firm for ease of notation. $IndicativeFee$ is the one day stock loan fee from Markit. $Realized\ Average\ Loan\ Fee_{t+m}$ is the one month realized average $IndicativeFee$ in month m. $MonthCost_{t,m}$ is the one month option shorting cost at time t in month m. $OptionPremium_{t,m}$ is the $MonthCost$ for month m minus the current $IndicativeFee$. Months $m = 1-6$ are shown. If markets are completely fungible and the expectations hypothesis holds, then the coefficient on $IndicativeFee$ and $OptionPremium$ should be 1. Robust standard errors double clustered by date and firm are shown below the parameter estimates in parenthesis. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively. I present p-values for the null where $OptionPremium_{t,m} = 1$.

Explanatory Variable	**Realized Average Loan Fee$_{t+m}$**					
Month	**1**	**2**	**3**	**4**	**5**	**6**
$OptionPremium_{t,m}$	0.041***	0.113***	0.194***	0.272***	0.225***	0.178***
	(0.010)	(0.028)	(0.042)	(0.056)	(0.059)	(0.054)
$IndicativeFee_t$	0.959***	0.909***	0.892***	0.894***	0.834***	0.776***
	(0.012)	(0.025)	(0.036)	(0.047)	(0.051)	(0.051)
Constant	-0.000	0.001*	0.001***	0.002***	0.003***	0.004***
	(0.000)	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)
Exp. hyp. p-value	0.000	0.000	0.000	0.000	0.000	0.000
N	60120	52131	51731	47464	44709	43683
R^2	0.921	0.740	0.597	0.504	0.431	0.393

Table 9

Option Shorting Cost Return Predictability Beyond the Current Stock Loan Fee

This table contains Fama-Macbeth (1973) results testing the relationship between option shorting costs, stock loan fees and returns. I suppress all i subscripts indicating firm for ease of notation. Ret_{t+m} is the buy and hold return in excess of the one-month risk-free rate in month m, $IndicativeFee$ is the one day stock loan fee from Markit, $MonthCost_{t,m}$ is the option shorting cost for month m and $OptionPremium_{t,m} = MonthCost_{t,m} - IndicativeFee_t$. $Book/Market$ is the log of the book/market ratio from Compustat, $MarketCap$ is the log of the market capitalization, $Idio.\ Volatility$ is the log of idiosyncratic volatility calculated using the monthly standard deviation of the residual from a Fama-French three-factor regression, $Bid - Ask$ is the log of the bid ask ratio and $Return_{t-1}$ is the return lagged by one month. If option shorting costs provide incremental return predictability beyond the current stock loan fee, then $OptionPremium$ should be negative. I report the time-series mean of the parameter estimates with t-statistics, calculated using Newey-West (1987) standard errors with three lags, shown below in parenthesis. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

Explanatory Variable	**Monthly Excess Return$_{t+m}$**											
Month	**1**	**1**	**2**	**2**	**3**	**3**	**4**	**4**	**5**	**5**	**6**	**6**
IndicativeFee$_t$	-1.416***	-1.280***	-1.280***	-1.075***	-1.187***	-0.756**	-1.095*	-1.055*	-2.019***	-1.426***	-1.100*	-0.688
	(-5.80)	(-3.80)	(-4.22)	(-3.30)	(-2.99)	(-2.08)	(-1.90)	(-1.93)	(-3.52)	(-3.07)	(-1.80)	(-1.17)
OptionPremium$_{t,m}$	-0.693***	-0.771**	-0.478	-0.506*	-0.757*	-0.523	-0.289	-0.415	-1.557**	-1.469**	-0.448	-0.064
	(-3.12)	(-2.58)	(-1.50)	(-1.71)	(-1.70)	(-1.22)	(-0.40)	(-0.60)	(-2.18)	(-2.39)	(-0.62)	(-0.09)
Book/Market		0.000		0.000		0.000		-0.000		-0.001		-0.001
		(0.11)		(0.18)		(0.20)		(-0.14)		(-0.71)		(-0.70)
MarketCap		0.001		0.000		0.001		0.001		-0.000		0.000
		(0.74)		(0.43)		(0.93)		(1.22)		(-0.29)		(0.41)
Idio. Volatility		-0.001		-0.006		0.003		-0.008		-0.018**		-0.000
		(-0.10)		(-0.96)		(0.52)		(-1.01)		(-2.33)		(-0.02)
Bid-Ask		0.000		0.000		-0.002		0.001		-0.002		0.001
		(0.20)		(0.30)		(-1.19)		(0.95)		(-1.29)		(0.93)
Return$_t$		-0.002		-0.007		-0.011		0.006		0.010		0.016
		(-0.19)		(-0.52)		(-1.01)		(0.77)		(0.83)		(1.52)
Return$_{t-1}$		-0.011		-0.001		0.010*		0.019		-0.003		0.012
		(-0.98)		(-0.08)		(1.69)		(1.52)		(-0.29)		(1.30)
Constant	0.008	-0.004	0.008	0.001	0.009	-0.022	0.007	-0.011	0.006	0.001	0.004	-0.003
	(1.24)	(-0.18)	(1.19)	(0.07)	(1.33)	(-1.00)	(1.06)	(-0.45)	(0.97)	(0.03)	(0.58)	(-0.12)
N	60517	47617	52751	41432	52494	41402	48270	38057	45647	36121	45100	35881
Average R^2	0.015	0.079	0.019	0.085	0.018	0.082	0.020	0.088	0.020	0.085	0.019	0.091

Table 10
The Cost Differential of Shorting in the Options Market and the Stock Loan Market and Recalls

This table contains results testing whether the difference between option shorting costs and the current stock loan fee predict recalls. I suppress all i subscripts indicating firm for ease of notation. $Recall_{t+m}$ is a dummy variable that equals 1 if shares available to be lent drops by 10% or more in month m. $IndicativeFee$ is the one day stock loan fee from Markit, $MonthCost_{t,m}$ is the option shorting cost for month m and $OptionPremium_{t,m} = MonthCost_{t,m} - IndicativeFee_t$. $MarketCap$ is the log of market capitalization, Ret_{t-1} is the return lagged by one month and $Volume_{t-1}$ is the total trading volume divided by shares outstanding, lagged by one month. Date fixed effects are used throughout and t-statistics are shown below the parameter estimates in parenthesis and are calculated using robust standard errors clustered by date. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

Explanatory Variable	**$Recall_{t+m}$**											
Month	**1**	**1**	**2**	**2**	**3**	**3**	**4**	**4**	**5**	**5**	**6**	**6**
$IndicativeFee_t$	0.283***	0.176***	0.337***	0.216***	0.379***	0.254***	0.443***	0.288***	0.456***	0.270***	0.477***	0.303***
	(7.88)	(4.80)	(8.25)	(4.93)	(8.73)	(5.78)	(7.87)	(5.19)	(6.73)	(4.39)	(8.27)	(5.46)
$OptionPremium_{t,m}$	0.138***	0.081***	0.239***	0.163***	0.255***	0.175***	0.302***	0.187***	0.334***	0.186***	0.302***	0.172**
	(5.57)	(3.38)	(6.29)	(4.13)	(4.99)	(3.54)	(4.92)	(3.05)	(4.74)	(2.89)	(4.35)	(2.51)
MarketCap		-0.010***		-0.011***		-0.011***		-0.011***		-0.012***		-0.011***
		(-7.62)		(-9.07)		(-9.21)		(-7.42)		(-7.74)		(-8.25)
$Volume_t$		0.022***		0.015*		0.029***		0.020***		0.014**		0.016**
		(3.09)		(1.88)		(4.44)		(2.80)		(2.31)		(2.09)
$Volume_{t-1}$		0.008		0.015**		0.002		0.005		0.014**		0.011
		(1.14)		(2.17)		(0.31)		(0.75)		(2.14)		(1.38)
$Return_t$		-0.018**		-0.012*		-0.040***		-0.030***		-0.018*		-0.032***
		(-2.25)		(-1.79)		(-4.04)		(-2.88)		(-1.87)		(-3.39)
$Return_{t-1}$		-0.006		-0.039***		-0.033***		-0.029***		-0.028***		-0.017*
		(-0.85)		(-3.74)		(-3.60)		(-2.84)		(-2.63)		(-1.79)
Constant	0.037***	0.245***	0.040***	0.273***	0.041***	0.283***	0.041***	0.282***	0.043***	0.296***	0.041***	0.284***
	(41.70)	(8.52)	(45.94)	(9.98)	(50.20)	(10.35)	(42.61)	(8.42)	(39.40)	(8.72)	(44.40)	(9.15)
Mean Dep. Var.	0.042	0.042	0.042	0.042	0.041	0.041	0.041	0.041	0.041	0.041	0.040	0.040
N	60622	60467	53003	52870	52942	52809	48936	48812	46548	46431	46229	46107
R^2	0.006	0.013	0.005	0.014	0.005	0.016	0.006	0.015	0.005	0.014	0.006	0.015

6.1 Appendix

Proposition 1: The assumption $E_t[P_{t+k}|t+k<\tau] = P_t$ yields the same equilibrium condition as if the price is constant until the information is revealed.

Proof: To calculate equilibrium short selling costs we need to find the expected payoff from shorting which is the current price minus the expected price at the end of the shorting contract, $E_t[P_t - P_{t+k}] = P_t - E_t[P_{t+k}]$. By the law of total expectation this equals:

$$= P_t - F_t(t+k)E_t[P_{t+k}|t+k \geq \tau] - (1 - F_t(t+k))E_t[P_{t+k}|t+k < \tau] \tag{12}$$

Since we assumed $E_t[P_{t+k}|t+k<\tau] = P_t$ and since $E_t[P_{t+k}|t+k \geq \tau] = V$ we have:

$$E_t[P_t - P_{t+k}] = F_t(t+k)(P_t - V) \tag{13}$$

Because of risk-neutrality and competitive equilibrium the expected payoff from shorting must equal the expected cost. Therefore, a per-period lending fee would be:

$$r_{t,k} = \frac{E_t[P_t - P_{t+k}]}{k} = \frac{F_t(t+k)(P_t - V)}{k} \tag{14}$$

Which is the same as in (2). ■