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Three Essays on Financial Markets and Institutional Investors

by

Blake Phillips

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Examining Committee

Aditya Kaul, Ph.D., Finance, University of Alberta

Vikas Mehrotra, Ph. D., Finance, University of Alberta

Susan Christoffersen, Ph.D., Finance, McGill University

Martin Luckert, Ph.D., Rural Economy, University of Alberta

David McLean, Ph.D., Finance, University of Alberta

Dedication

To my beloved wife and children, for without your love and support this would never have been achievable.

Abstract

Chapter 2 undertakes a new investigation of the potential for options to mitigate short sale constraints, conducing two event studies which examine 1732 option introductions and the differential effect of the 2008 short sale ban on optioned and non-optioned stocks. I find option introduction mitigates 79% of the price adjustment efficiency disparity between short sale constrained and unconstrained stocks in relation to negative news. I also find evidence that negative information was incorporated more freely into optioned stocks during the short sale transaction ban of financial sector stocks. These results collectively suggest that in the presence of binding short sale constraints, options act as an effective substitute to short sales, significantly contributing to the informational efficiency of the market.

In Chapter 3 we examine the determinants of success of foreign cross-listings in the U.S. using cumulative returns surrounding the cross-listing event and liquidity on the U.S. exchange as joint metrics of success. We find that the post-listing liquidity and valuation benefits of cross-listings are crucially dependent both on prior home-market success and on U.S. institutional holdings in the cross-listing quarter. Stocks with greater institutional ownership upon cross-listing see more liquid U.S. trading. Additionally, firms with a higher abnormal price run-up in the year prior to cross-listing and firms that see more liquid domestic trading enjoy greater post-listing liquidity in the U.S.

Chapter 4 examines the asset allocation decisions of mutual fund investors, focusing on flight to quality considerations. Using the default spread, term spread and short term interest rate as proxies for economic conditions, we find that an expected improvement (deterioration) in Canadian economic conditions causes investors to direct flow away from (towards) fixed income-type funds and towards (out of) equity based funds. For example, a one standard deviation increase in the term spread (1.13%) results in an 84% increase and a 74% decrease in the percentage of flow directed at Canadian equity and money market funds respectively, relative to the previous month.

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Chapter 1

Introduction

The U.S. Securities and Exchange Commission defines an institutional investor as a money manager with greater than $100 million at their discretion. These entities typically include banks, insurance companies, mutual funds, pension funds and hedge funds which act to pool the assets of individual investors and invest and manage those assets in aggregate. In 2005 the International Monetary Fund estimated that over 45 trillion dollars in assets were under management by institutional investors throughout the world (IMF, 2005).

The role and effects of institutional investors in financial markets has been examined extensively in the finance literature. As institutional investors have large stakes in the performance of firms in which they invest (relative to small individual investors), they have greater motivation to undertake costly information seeking. Additionally, economies of scale potentially allow institutions to undertake information collection at reduced costs. In support of theories of enhanced price discovery stemming from institutional investors, Boehmer and Kelley (2007) find stocks with greater institutional investor holdings are priced more efficiently.

In this thesis I further examine the role of institutions in financial markets, focusing predominantly on their role in information collection and price efficiency. First, in chapter 2, I examine the importance of institutional investors as providers of short sale liquidity. Institutions in large, long positions (such as pension or index funds) provide the majority of liquidity for short sale transactions (D'Avolio, 2002). In the absence of short sales, stock price adjustment efficiency becomes delayed (Diamond and Verrecchia, 1987) and pessimistic investors who do not own the stock are excluded from investing on their value opinions (Miller, 1977). The joint effect is an upward bias in prices resulting from the exclusion of negative sentiment and information from the market.

Chapter 2 examines the potential for options to mitigate short sale constraints stemming from low or absent short sale loan supply from institutional investors. Once option trading is introduced, investors may realize a synthetic, short position, circumventing short sale constraints. The findings of this chapter point to the critical role which institutions play in enabling short sale transactions and the importance of short sales for efficient markets.

Chapter 3 focuses on the determinants of success of foreign cross-listing in the U.S. We use cumulative returns the year before and after cross-listing and liquidity on the U.S. exchange as twin metrics of cross-listing success. We also examine the role which U.S. institutional investors play in cross-listings. Given the importance of institutional investors for efficient information seeking and price discovery, we would expect the support of institutions to be critical to cross-listing success. In consideration of the perception of institutions as informed investors, high institutional ownership at the time of cross-listing could be viewed as a signal of quality to other investors. Further, institutions with large holdings would potentially contribute to increased price pressure and enhanced liquidity, contributing positively to our key measures of cross-listing success. Consistent with this view, we find institutional holdings are linked to sustaining pre cross-listing price run-ups and greater liquidity on the U.S. exchange, both factors noted as objectives of cross-listing by firm managers (Mittoo, 1992).

In Chapter 4 we deviate from our examination of the role of institutional investors in the market and instead examine the effect that macroeconomic factors have on the flow of investment assets to mutual funds. The success of mutual fund managers depends in part on their ability to anticipate asset flows and develop an effective investment strategy around that asset base. Coval and Stafford (2007) show that mutual funds that experience high inflows or outflows of net assets significantly under-perform funds with more moderate asset flows. Using bond spreads as predictors of future economic performance we show investors vary asset allocations across broad asset risk classes in anticipation of changing macroeconomic conditions. We then examine whether this switching strategy is potentially beneficial to investors from a risk-adjusted return perspective.

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Chapter 2

Options, Short Sale Constraints And Market Efficiency

"Short sellers occupy a position in the stock market like that *of predators in nature: necessary but unloved" (Sauer, 2006).*

Through history short sellers have been both reviled and lauded by investors. Short selling "bear raids" were widely perceived by many investors as a cause of the 1929 stock market crash (U.S. SEC, 2007). More recently, on September 19, 2008 the U.S. Securities and Exchange Commission (SEC) banned short sale transactions for financial sector stocks in an effort to stabilize the market amidst the 2008 global financial crisis. The effect of short sale constraints on stock prices and market efficiency has been studied extensively. Most empirical studies conclude that short sale constraints are associated with overpricing and are related to a reduction in market quality and the efficiency which negative information incorporates into stock prices.[1]

The finance literature is less clear and divided on the potential for options to mitigate short sale constraints. Several authors have found evidence in support of a reduction in short sale constraints following option introduction.[2] These results have been criticized by Mayhew and Mihov (2005) for failing to control for option introduction endogeneity, suggesting that stock characteristics previously regarded as evidence of a reduction of short sale constraints may be spurious. Further, Bris et

[1] For example D'Avolio (2002) and Nagel (2005) both conclude short sale constraints contribute to the exclusion of negative value opinions from the market leading to overpricing. Charoenrook and Daouk, (2005) analyze markets in 111 countries and find that when short selling is allowed expected returns are lower and that stock prices exhibit less volatility and greater liquidity. Bris et al. (2007) examine 46 markets around the world and conclude negative information is incorporated into stock prices faster when short selling is allowed and practiced.

[2] For example, Figlewski and Webb (1993) find that short interest increases and Sorescu (2000) finds negative abnormal returns for stocks following option introduction. Both findings are reflective of greater access to synthetic short transactions and the correction of overpricing by options. This literature is discussed in greater detail in Section I.

al. (2007) examine forty-six equity markets around the world and find the effect of put options to be insignificant in the presence of short selling restrictions. In this paper I contribute to this ongoing debate by conducting two event studies, the first examines the change in market efficiency following 1732 option introduction events between 1981 and 1997. The second examines differences in cumulative stock returns for optioned and non-optioned stocks during the short sale ban from September 19 to October 8, 2008. [3]

This paper makes several important contributions to the finance literature. First, this is the first paper to examine the change in market efficiency resulting from option mitigation of short sale constraints which controls for option introduction endogeneity. Second, Asquith et al. (2005) and others argue short sale constraints likely only bind for a minority of stocks for which short sale demand exceeds share loan supply. Thus, prior literature which examines the marginal effect of option introduction on market short sale constraints, may fail to detect a significant effect for a noteworthy subset of the market. To my knowledge, this is the first article to examine cross-sectional variation in the change in market efficiency across short sale constraint levels documenting a direct link between options and short sale constraint reduction. Finally, also to my knowledge, this is the first paper to utilize the natural experiment of the 2008 short sale ban to further examine the potential for options to mitigate short sale constraints.

There are two prevalent hypotheses in relation to the source of short sale constraint effects. Miller (1977) argues that short sale constraints exclude

[3] The option introduction dataset is as used by Mayhew and Mihov (2004).

pessimistic investors from the market, thus when short sale constraints bind, stock prices are set by the subset of investors with the most optimistic value opinion of the stock. In slight contrast, Diamond and Verrecchia (1987) argue that once option trading is introduced, investors may take a synthetic short position allowing more rapid incorporation of negative, private information into stock prices. As an extension of the market efficiency tests, I examine the differential improvement in market efficiency in relation to private and public news. These additional tests allow inferences regarding the source of the short sale constraint effect, contrasting the behavioral source suggested by Miller (1977) to the rational movement of markets from semi-strong to strong form efficiency suggested by Diamond and Verrecchia (1987).

In the option introduction event study, to measure post-option reduction in short sale constraints I utilize an extension of the adjustment delay measures defined in Hou and Moskowitz (2005).[4] The Hou and Moskowitz model assesses the significance of lagged market returns for predicting contemporaneous stock returns. The greater the number of lagged market returns that are significant for predicting contemporaneous stock returns, the greater the delay in adjustment to new information.[5] Similar to Danielsen et al. (2007) and others, I find that on average short sale constraints don't bind prior to option introduction. Only stocks

[4] The Hou and Moskowitz (2005) model is similarly used to quantify the extent of short sale constraints by Saffi and Sigurdsson (2007) and is similar to the model utilized by Bris et al. (2007).

[5] As discussed in more detail in Section III, I augment the original Hou and Moskowitz (2005) model with excess stock returns and negative news interaction dummies to allow a greater range of market efficiency tests.

with low institutional ownership, with low short sale loan supply, realize a significant post-option improvement in the speed of price adjustment.[6]

Consistent with the hypothesis that post-option improvement in market efficiency is related to the relaxation of short sale constraints, I find that the improvement is limited to negative news. No significant post-option improvement in adjustment efficiency is noted in response to positive news. To my knowledge this is the first paper to report this result. Prior to option introduction, short sale constrained stocks adjust to negative news 19% more slowly than unconstrained stocks. Following option introduction that difference is reduced to 4%, indicating that options eliminate up to 79% of the price efficiency disparity between short sale constrained and unconstrained stocks. These results are robust to the control methodology suggested by the results of Danielsen et al. (2007).

As an extension of these tests I replicate the extended Hou and Moskowitz models utilizing lagged excess stock returns in the place of lagged market returns. As market returns reflect predominantly public information and excess stock returns reflect both private and public firm specific information, the Diamond and Verrecchia (1987) and Miller (1977) hypotheses generate contrasting predictions. Following Diamond and Verrecchia, efficiency gains should be limited to the negative, private information component of firm specific returns. In contrast, Miller's hypothesis predicts similar efficiency gains in relation to the public information component of both market and excess stock returns. I find that improvements in adjustment efficiency are similar between the market and excess

[6] To proxy for short sale loan supply I use institutional ownership, as institutional investors in long positions provide the majority of shares for short sale loans (D'Avolio, 2002)

stock return models, suggesting short sale constraint effects stem from an optimism bias or other behavioral source as argued by Miller (1977). Delayed incorporation of negative, private information into stock prices likely also plays a role in short sale constraint effects, but does not appear to be the exclusive source as suggested by Diamond and Verrecchia (1987).

As a final test of the potential for options to mitigate short sale constraints I examine cumulative returns to optioned and non-optioned stocks during the 2008 short sale ban. As option exchange bookmakers were excluded from the short sale ban, they were able to continue to hedge their put option exposure via short sales, allowing option markets to remain functional. During the short sale ban from September 22 to October 8, 2008, banned optioned and non-optioned stock prices dropped on average 29% and 16%, respectively. The difference of 13% is statistically significant and is robust to controls for differences in stock characteristic and sensitivity to financial sector news between the two sub-samples. This result suggests negative information was incorporated more freely into optioned stock prices and that options acted as a substitute to short sales during the short sale ban.

Considered collectively the results of this paper can be summarized as follows. For the majority of the market, adequate short sale loan supply is available to meet demand such that short sale constraints do not bind, even in the absence of options. For a subset of the market with low institutional investor holdings, who typically provide the majority of short sale liquidity, options act as a substitute for short sales contributing significantly to market efficiency. On occasions when short sale constraints are artificially imposed via legislative action, options also serve to mitigate short sale constraints, potentially undermining the objective and intent of SEC market oversight actions in September 2008.

The remainder of this chapter is organized as follows: Section I reviews the related literature in more detail. Section II describes the sample and Section III describes the stock price adjustment delay measures utilized in the chapter. Section IV presents an analysis of the determinants of stock price delay. Section V and VI examine the effect of option introduction on the speed of stock price adjustment to market-wide and firm specific information, respectively. Section VII presents robustness checks for the option introduction analysis, Section VIII evaluates optioned and non-optioned stocks during the short sale ban and Section IX concludes the chapter.

I. Related Literature

This chapter relates to two bodies of research: (1) the effect of option introduction on short sale constraints and (2) the effect of options on price adjustment of the underlying stock. Subsection A reviews the extant empirical results related to option introduction and short sale constraints. Subsection B describes the process by which option introduction may affect the price adjustment efficiency of the underlying stock.

a. *Option Introduction and Short Sale Constraints*

Empirical work testing the effect of option introduction on short sale constraints can be characterized as following four different approaches. First, several authors have investigated the change in short interest following option introduction. If option book makers face lower short selling constraints than the average investor, option introduction may result in increased short interest as option bookmakers hedge synthetic short positions taken by individual investors in the option market. In support of this hypothesis, Damodaran and Lim (1992), Danielsen and Sorescu (2001) and Figlewski and Webb (1993) all find short interest increases following option introduction.

Second, Mayhew and Mihov (2005) investigate option trading volume following option introduction. After examining nine proxies for short sale constraints, either no relationship or a significantly negative relationship between short sale constraint proxies and option trading volume was found. While this finding lends little support to the Diamond and Verrecchia (1987) hypothesis, it is possible that the option market contributes to the price discovery of the underlying stock via other mechanisms (other than trading volume). For example, Chan et al. (2002) examine the intraday interdependence of order flows and price movements for actively traded New York Stock Exchange (NYSE) stocks and Chicago Board Options Exchange (CBOE) traded options. Their findings indicate that while information in the stock market is contained in both quote revisions and in net trade volume, information in the option market is contained only in quote revisions. [7]

Third, researchers have examined return behavior following option introduction. If option introduction reduces short sale constraints, it would be expected to be predictive of negative abnormal returns as historical negative information withheld from the market is impounded in stock prices. Using the value weighted market index as a reference portfolio, Sorescu (2000) and Danielsen and Sorescu (2001) find that for options listed from 1980 to 1995 the underlying stock experiences negative abnormal returns following option introduction. Mayhew and Mihov (2005) undertake the same analysis using a control sample of non-optioned stocks which have similar characteristics to those selected for option introduction. The control portfolio is found to exhibit similar negative abnormal returns, suggesting that the relationship between option introduction and negative abnormal returns may be spurious and more a result of stock characteristics common at the time of option listing.

Fourth, the approach of this paper and other authors has been to investigate the efficiency by which negative and positive information is incorporated into stock prices following option introduction.[8] Most closely related to this paper, Damodaran and Lim (1992) examine stock return processes following option introduction focusing on mean reversion and skewness. Using a relatively small sample of 200 firms with option introduction from 1977-1984, they conduct an

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[7] Net trading volume is defined as buyer initiated trading volume minus seller initiated trading volume.

[8] See for example Jennings and Starks (1986), Skinner (1990) and Damodaran and Lim (1991)

event study of cumulative abnormal returns over the 20 days surrounding earnings announcements before and after option introduction. They find that following option introduction a greater amount of the information related to earnings announcement shocks is impounded in stock prices in the ten days prior to earnings announcements and that prices adjust more quickly to negative earnings shocks after option introduction. They interpret these results as supportive of the hypothesis that easing of short sale constraints allows stock prices to adjust more rapidly to negative information.

As shown by Mayhew and Mihov (2005), a viable alternative hypothesis is that endogenous stock characteristics common at the time of option introduction contribute to faster incorporation of negative information into stock prices. I extend the work of Damodaran and Lim (1992) by controlling for endogenous stock characteristics utilizing the control methodology specified by Danielsen et al. (2007). Further, I utilize a much larger sample (1732 firms) over a greater time frame (1981 – 1997) to test for a direct relation between improved post option price adjustment and short sale constraints. The unique stock price adjustment model utilized in this paper allows the comparison of stock characteristics the year before and after option introduction. Due to sample size and methodology limitations, Damodaran and Lim (1992) use 10 years of pre option listing data which further exposes the model to potential endogeneity issues.

b. Option Introduction and the Speed of Stock Price Adjustment

Option introduction has the potential to improve the speed of stock price adjustment to new information via four unique channels. First, on average, option introduction reduces the bid-ask spread of the underlying equity (Fedenia and Grammatikos, 1992), thereby reducing transaction costs and the magnitude of disparity between stock price and the perceived value necessary to trigger trading. Thus, smaller magnitude information events should be incorporated into stock prices more rapidly following option introduction.

Second, the option exchange provides an alternative venue by which private information and pessimistic opinions may be made public and contribute to price discovery of the underlying stock. In a general context, investors with access to private information can choose to trade on either the stock or option market. On average, Chakravarty et al. (2004) find the option market contribution to price discovery to be approximately 17%. In the context of option introduction for a short sale constrained stock, informed investors with negative information who wish to short the stock are forced to trade on the option market. Thus, the expected contribution of the option market in that context would likely be higher.

Third, the addition of option trading and the associated ease of hedging positions in the underlying stock may attract greater institutional ownership. To the extent which institutional investors are more informed and sophisticated than the average investor, greater institutional investor holdings could contribute to improved informational efficiency.

Finally, as pessimistic investors take synthetic short positions on the option market, the option book maker presumably hedges his exposure via short selling, placing downward price pressure on the underlying stock. The obvious question becomes how the option book maker is able to hedge his position via short sales, when the pessimistic investor is not? For several reasons, the option book maker has access to short sale loan supply not available to the average investor.

First, transaction costs related to short selling can vary across agents depending on trading volume and frequency. In a typical short sale contract, the short seller borrows the shares from his broker and the proceeds of the sale guarantee the loan and generate interest. The broker returns a portion of the interest to the short seller at the conclusion of the contract, known as the rebate rate. Where loan demand exceeds loan supply, rebate rates may become negative, indicating the borrower pays a fee to the lender for the opportunity to borrow the stock. In support of Miller's 1977 theory, stocks with negative rebate rates (high loan demand) tend to under-perform in the future, indicating overpricing (Jones and Lamont, 2002). Brokers vary rebate rates and share availability preferentially across clients, with high volume customers receiving more favorable treatment (Evans et al., 2003). Thus, short sales for low loan supply (negative rebate rate) stocks may be impossible or prohibitively costly for the average investor relative to high volume borrowers, such as option book makers.

Second, exchange rules require most market participants to demonstrate they are able to borrow low rebate rate stocks prior to short selling (Evans et al., 2003). Market book makers are exempt from this requirement by NASD rule 3370. If unable to locate shares following initiation of the short sale, the bookmaker has the additional option to fail to deliver. Once the bookmaker fails to deliver, one of two scenarios occur. The buyer's broker may allow the failure to deliver to continue as long as the short sale contract is open. In this situation the consequence of failing to deliver is forgoing the interest on the proceeds of the sale but the short position is effectively maintained without delivery (Evans et al., 2003). Alternatively, the buyer's broker may insist on delivery and file a notice of intention to buy-in, in which case the short seller has two days to deliver the shares or the buyer purchases the shares on the short seller's account (Evans et al., 2003). In the event of a buy-in, the market maker must short sell again to re-establish the position, incurring execution costs plus the difference between the buy-in and market price. Utilizing a two year database of short sale transaction data for 1998 and 1999, Evans et al. (2003) find buy-ins occur in only 0.12% of failures to deliver. Thus, exchange rules give option bookmakers a regulatory advantage which allows them to short sell without actually borrowing stock.

II. Sample Description

Included in the option introduction sample are all stocks for which an option was introduced from January 1981 through January 1997 as used in Mayhew and Mihov (2004). The original Mayhew and Mihov dataset was provided by the Chicago Board Option Exchange (CBOE) and includes all option introductions on the CBOE, the American, Philadelphia, Midwest and Pacific Exchanges from January 1973 – January 1997.

As a proxy for short sale constraints I use institutional ownership measured in terms of percentage of shares outstanding held.[9] Intuitively, investors in large, long positions would be best suited to fulfill the supply side of short sale loan transactions. In support of this hypothesis, D'Avolio (2002) documents that institutional investors provide the majority of stock loan supply for short selling.[10] Institutional ownership data are obtained from the Shareworld 13F Filing database as maintained by Thomson Financial. The Shareworld 13F database tracks, by quarter, the share holdings of institutional investors based on 13F filings made with the U.S. Securities and Exchange Commission. Any manager with more than $100 million at their discretion is required to make a quarterly filing of a 13F for every security holding in excess of $200,000 or 10,000 shares. Additionally, the Shareworld 13F database tracks ownership profiles of non-U.S. equities based on the Information Sheets and Shareholder Reports of both Domestic and Foreign Mutual Funds. The institutions represented in Shareworld include mutual funds, banks, insurance firms, and pension funds. At the time of dataset construction, the Shareworld database tracked 13F filings from 1980 through 2005. As stock and

[9] Results are robust to an alternative specification using the number of institutional owners instead of shares held.

[10] Following research by Nagel (2005), Chen et al. (2002) and others it is widely accepted that institutional ownership is an effective proxy for short sale constraints. Further evidence in support of the use of institutional ownership as a proxy for short sale constraints is provided by Asquith et al. (2005) who document that the percentage of shares outstanding shorted is less than the percentage of shares held by institutions in 95% of their dataset. Asquith et al. (2005) suggest an alternative short sale constraint proxy defined as the interaction of low short sale loan supply (low institutional ownership) and high short sale demand (high short interest). Due to data availability limitations I am unable to include short interest as a variable in this chapter. This limitation, if anything, biases against the findings of the chapter as stocks classified as unconstrained (moderate to high institutional ownership) are potentially constrained if short sale demand exceeds supply.

return characteristics one year before and after option introduction are desired, all option introductions prior to 1981 are excluded. Additionally, stocks for which the option de-listed within one year of introduction are also excluded (108 stocks).

Figure 1 presents mean quarterly institutional ownership for the option introduction sample one year before and after option introduction. Through the four quarters preceding option introduction, mean institutional ownership steadily increases from 27% to 33% of total shares outstanding. The average number of institutional investors holding shares in each stock experiences a similar rate of increase over the same timeframe.

Stocks under consideration for option listing are known only to the members of the board of the option exchange and option introduction announcements are made public on average 3 trading days prior to the initiation of trading (Danielsen et al., 2007). Given the timing of option introduction announcements, the increase in institutional ownership prior to option introduction does not reflect a desire for institutions to hold stocks that eventually have traded options. Mayhew and Mihov (2004) document that option exchanges tend to select stocks with high volatility, trading volume and name recognition when determining new listings. The increase in institutional ownership preceding option introduction reflects that institutional investors are attracted to a similar set of stock characteristics.

Table I presents summary statistics for the option introduction sample at the time of option introduction. The mean and median stock prices at the time of option introduction are $26.45 and $23.25 respectively.[11] On average, stocks in the test sample trade 120,000 shares per day, which represents a turnover of 0.7% of total shares outstanding (annualized turnover ratio of 1.75). Mean market capitalization is $805 million but the mean is biased upwards by several high market capitalization stocks, reflected by the median market capitalization value of $470 million.

The short sale ban sample includes all stocks which were placed on the SEC short sale ban list between September 19, 2008 and September 26, 2008. Excluded from the sample are all stocks that were subsequently removed from the banned list at the request of the firm (10 stocks) and stocks with less than 240 return observations over the year prior to short sale ban. In total 986 stocks were placed on the banned list, 916 met the inclusion criteria, 88% of which were placed on the list on September 19, 2008, 10% of which were added on September 22 or 23, 2008. Stock price, volume and shares outstanding data were obtained from DataStream for the short sale ban sample.

Table II presents summary statistics for the short sale ban sample, summarized separately for optioned and non-optioned stocks. Consistent with the intuition suggested by Mayhew and Mihov (2004), on average optioned stocks tend to be larger, with greater shares outstanding and experience higher trading frequency as measured by daily average turnover. The average price of the non-

[11] Stock price, volume and shares outstanding data were obtained from the Center for Research in Security Prices database.

optioned sample is skewed significantly upwards by the inclusion of Berkshire Hathaway A and B class stocks.

III. Speed of Stock Price Adjustment Measures

To measure the speed which new information is impounded in stock prices I utilize the stock price adjustment delay measures developed by Hou and Moskowitz (2005). In this model, the market return is utilized as a proxy for the new public information to which individual stock prices respond. [12]

$$r_{j,t} = \alpha_j + \beta_j^0 R_{m,t} + \varepsilon_{j,t} \qquad \textbf{(\textit{base model})} \qquad (2.1)$$

$$r_{j,t} = \alpha_j + \beta_j^0 R_{m,t} + \sum_{n=1}^{5} \beta_j^n R_{m,t-n} + \varepsilon_{j,t} \qquad \textbf{(\textit{extended market model})} \qquad (2.2)$$

Here $r_{j,t}$ is the return of stock *j* on day *t*, $R_{m,t}$ is the market return on day *t* and $R_{m,t-n}$ is the market return *n* days prior to day *t*. In the extended market model, if the stock responds immediately to new information, β_j^0 will be significantly different from zero, but none of the β_j^n will differ significantly from zero. On the other hand, if the response is delayed β_j^0 will be less significant or insignificant and some or all of the β_j^n will be significantly different than zero.

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[12] The Equal Weighted Index Return (excluding distributions), as maintained on the Center for Research in Security Prices database, was utilized as a proxy for market return.

Using Akaike's and Schwarz's information criterion (AIC and SIC respectively) the goodness of fit of the extended model is optimized with the inclusion of five lags for 77% (using AIC) and 96% (using SIC) of the test sample stocks. Further optimization of AIC and SIC values for the remainder of the stocks requires the inclusion of 6 or more additional lags. Five lags are selected to optimize the AIC and SIC values for the majority of stocks in the sample.

The first delay measure used by Hou and Moskowitz (2005), the R^2 Ratio, measures the proportional difference between the explanatory power of contemporaneous versus lagged market returns to predict stock returns.

$$D_{rsq} = 1 - \frac{R^2_{base}}{R^2_{extended}} \tag{2.3}$$

The faster new information is incorporated into individual stock prices, the smaller the difference between the R^2 of the extended and base models, as lagged market returns add little by the way of explanatory power. Thus, as the speed of stock price adjustment increases the D_{rsq} delay measure decreases.

The second delay measure, the Coefficient Ratio, measures the ratio of the lag-weighted sum of the lagged market return coefficients relative to the sum of all the regression coefficients. Similar to the R^2 Ratio delay measure, the greater the delay in stock price adjustment, the larger the lagged regression coefficients and the larger the D_{sum} delay measure.

$$D_{sum} = \frac{\sum_{n=1}^{5} n(abs(\beta_j^n))}{abs(\beta_j^0) + \sum_{n=1}^{5} abs(\beta_j^n)} \qquad (2.4)$$

The third delay measure, the Standard Error Adjusted Coefficient Ratio, augments the Coefficient Ratio measure by weighting each coefficient by its standard error. Thus, the significance of each coefficient is considered, whereas the raw Coefficient Ratio only considers the magnitude of the regression coefficients.

$$D_{se} = \frac{\sum_{n=1}^{5} \frac{n(abs(\beta_j^n))}{se(\beta_j^n)}}{\frac{abs(\beta_j^0)}{se(\beta_j^0)} + \sum_{n=1}^{5} \frac{abs(\beta_j^n)}{se(\beta_j^n)}} \qquad (2.5)$$

To capture the speed of stock price adjustment to firm specific information the same methodology is utilized but lagged excess stock returns are used in the place of lagged market returns.

$$r_{j,t} = \alpha_j + \beta_j^0 R_{m,t} + \sum_{n=1}^{5} \beta_j^n R_{j,t-n} + \varepsilon_{j,t} \quad \textbf{(\textit{extended firm model})} \qquad (2.6)$$

Here $r_{j,t}$ is the return of stock *j* on day *t*, $R_{m,t}$ is the market return on day *t* and $R_{j,t-n}$ is the return to stock *j* *n* days prior to day *t* less the market return *n* days prior to day *t*. The base model of comparison remains unchanged and the delay measures are calculated in the same manner, contrasting the R^2 and coefficient values of the base model and the extended firm model. Throughout the paper I use DM to designate delay measures derived utilizing lagged market returns and DF to designate delay measures derived utilizing lagged excess stock returns.

To allow separate quantification of the speed of adjustment to negative and positive news, I augment the extended models with negative news interaction dummy variables.

(extended neg market model)

$$r_{j,t} = \alpha_j + \beta_j^0 R_{m,t} + \beta_j^{-d0} D_j^0 R_{m,t} + \sum_{n=1}^{5} \beta_j^n R_{m,t-n} + \sum_{n=1}^{5} \beta_j^{-dn} D_j^n R_{m,t-n} + \varepsilon_{j,t} \qquad (2.7)$$

(extended neg firm model)

$$r_{j,t} = \alpha_j + \beta_j^0 R_{m,t} + \beta_j^{-d0} D_j^0 R_{m,t} + \sum_{n=1}^{5} \beta_j^n R_{j,t-n} + \sum_{n=1}^{5} \beta_j^{-dn} D_j^n R_{j,t-n} + \varepsilon_{j,t} \qquad (2.8)$$

D_j^0 is set to 1 if the contemporaneous market return is negative, otherwise it is equal to zero. Likewise, in the extended market model each D_j^n is equal to 1 if the market return for that specific lag is negative, otherwise it is equal to zero. Similarly, in the extended firm model each D_j^n is equal to 1 if the excess stock return for that specific lag is negative, otherwise it is equal to zero. Within the extended neg models, the β_j^n coefficients reflect the relation between stock returns and lagged positive market or excess stock returns while the β_j^{-dn} coefficients reflect the incremental effects of lagged negative market or excess stock returns.

D_{rsq}^{neg} is calculated as the R^2 Ratio of the extended neg model relative to the extended model.

$$D_{rsq}^{neg} = 1 - \frac{R_{extended}^2}{R_{extended-neg}^2} \tag{2.9}$$

Higher values of D_{rsq}^{neg} reflect a greater delay in the speed of price adjustment to negative new information. D_{sum}^{neg} and D_{se}^{neg}, which contrast the β_j^{-dn} coefficients to β_j^0 in the extended neg model, are additional measures of the price delay related to negative news:

$$D_{sum}^{neg} = \frac{\sum_{n=1}^{5} n(abs(\beta_j^{-dn}))}{abs(\beta_j^0) + abs(\beta_j^{-d0}) + \sum_{n=1}^{5} abs(\beta_j^n) + \sum_{n=1}^{5} abs(\beta_j^{-dn})} \tag{2.10}$$

$$D_{se}^{neg} = \frac{\sum_{n=1}^{5} \frac{n(abs(\beta_j^{-dn}))}{se(\beta_j^{-dn})}}{\frac{abs(\beta_j^0)}{se(\beta_j^0)} + \frac{abs(\beta_j^{-d0})}{se(\beta_j^{-d0})} + \sum_{n=1}^{5} \frac{abs(\beta_j^n)}{se(\beta_j^n)} + \sum_{n=1}^{5} \frac{abs(\beta_j^{-dn})}{se(\beta_j^{-dn})}} \tag{2.11}$$

To quantify the speed of stock price adjustment related to positive news D_{sum}^{pos} and D_{se}^{pos} are calculated as the ratio between the β_j^n coefficients relative to the β_j^0 coefficient in the extended neg model. However, given that the positive and negative interaction effects have an identical effect on R^2, D_{rsq}^{pos} is non-informative in this context.

$$D_{sum}^{pos} = \frac{\sum_{n=1}^{5} n(abs(\beta_j^n))}{abs(\beta_j^0) + \sum_{n=1}^{5} abs(\beta_j^n)} \tag{2.12}$$

$$D_{se}^{pos} = \frac{\sum_{n=1}^{5} \frac{n(abs(\beta_j^n))}{se(\beta_j^n)}}{\frac{abs(\beta_j^0)}{se(\beta_j^0)} + \sum_{n=1}^{5} \frac{abs(\beta_j^n)}{se(\beta_j^n)}} \tag{2.13}$$

IV. Determinants of Stock Price Adjustment Delay

Due to the short duration of the short sale ban (13 trading days) the speed of stock price adjustment analysis is conducted only for the option introduction sample. A detailed analysis of the determinants of stock price adjustment delay is completed by Hou and Moskowitz (2005). Thus, I provide only a brief analysis to provide a sense of the characteristics of the delay measures, demonstrate consistency of the extended delay measures with the extant literature and to provide motivation for the control variables utilized further in the paper.

Potential determinants of stock price adjustment fall into two categories: first, adjustment may be delayed due to microstructure factors such as a small investor base, low liquidity or limited information arrival. As microstructure determinants of stock price adjustment efficiency I use turnover (TURN), illiquidity (ILLIQ), market capitalization (SIZE) and the standard deviation of stock returns (VOL). Second, stock price adjustment may be delayed due to investor inattention, thus as proxies for investor inattention I include institutional ownership (INST), book to market (BK/MK), number of employees (EMPLOY) and number of analysts (ANAL).[13] INST is the percentage of total shares outstanding held in aggregate by institutional investors in the option introduction quarter. BK/MK is book value in the year of option introduction divided by market capitalization the day before option introduction.[14] TURN is the mean daily turnover the year preceding option introduction.[15] ILLIQ is the average weekly Amihud Illiquidity Ratio (Amihud, 2002) the year preceding option introduction:

$$IR_d = \frac{|r_d|}{Volume_d} \tag{2.14}$$

[13] The selected determinants of stock price adjustment are similar to those used in Hou and Moskowitz (2005) and variables found to influence short sale constraint levels in Nagel (2005). They are also consistent with variables found by Mayhew and Mihov (2004) and Danielsen et al. (2007) to predict option introduction likelihood.

[14] Book value data was obtained from the Compustat database, where book value is defined as common equity plus balance sheet deferred taxes. Book value data could not be located for 13 stocks. For those stocks, book value / market value was set to the average of the test sample. Findings are robust if those 13 stocks are excluded from the sample.

[15] Turnover is calculated as daily trading volume divided by shares outstanding.

where r_d is the close-close weekly return and $Volume_d$ is the dollar value of aggregate weekly volume, both in week d.[16] SIZE is market capitalization the day prior to option introduction. VOL is the standard deviation of daily stock returns over the year preceding option introduction. ANAL is the number of analysts which cover the stock in the year of option introduction and EMPLOY is the number of people employed by the firm in the fiscal year end of option introduction.[17]

Table III presents the correlation matrix for the determinant variables of stock price adjustment delay. Correlation levels are generally below 0.30, with the exception of the correlation between return volatility and turnover (0.56) and between return volatility and size (-0.48). These correlations are intuitive: high volatility stocks tend to be small (low market capitalization) with high turnover.

Table IV reports the cross-sectional regression results of the stock price delay measures regressed on the selected determinant variables, where the stock price adjustment delay measures are estimated over the year prior to option introduction. As the correlation between the D_{sum} and D_{se} delay measures is in excess of 0.9 in both the pre and post option periods across the majority of models, in the interest of brevity, only D_{rsq} and D_{se} values are reported for each model throughout the chapter. The D_{sum} values are available upon request. There is a high level of consistency in the determinant regression results, where institutional ownership, size and

[16] Weekly returns and aggregate weekly volume are used as opposed to daily values to control for downward bias in the Amihud illiquidity measure which potentially result from thin trading. Results are robust if daily volume and return data are used instead.

[17] Employment data was obtained from the Compustat database and analyst coverage data was obtained from the International Brokers Estimates System database. Employment data was unavailable for approximately 10% of the test sample, thus for those stocks the mean employment value of the sample was used.

illiquidity are significant determinants of stock price adjustment delay in the majority of the models. Analyst coverage and stock price volatility are significant in approximately half the models, while book to market, turnover and employment are rarely significant. Hou and Moskowitz (2005) document that high delay firms tend to be small, volatile and less visible stocks potentially overlooked or neglected by investors. Intuitively, stocks for which prices adjust rapidly will be characterized by a large, active investor base which is able to quickly evaluate and trade on new information and vice versa. I find results consistent with this conjecture. In the pre option introduction period, across both speed of adjustment measures (D_{rsq} and D_{se}), large stocks with high institutional ownership, analyst coverage, number of employees and turnover (negative coefficient values for INST, TURN, ANAL, EMPLOY and SIZE) and low volatility and illiquidity (positive coefficient values for VOL and ILLIQ) tend to be stocks with high price adjustment efficiency (low D_{rsq} and D_{se} values).[18]

V. Option Introduction and Stock Price Adjustment to Market News

The first question of my analysis examines whether option introduction contributes to the relaxation of short sale constraints. I first examine the post option change in market efficiency in relation to public, market-wide news proxied by daily market returns as this model construct is most consistent with Hou and Moskowitz (2005). Results related to the change in market efficiency in relation to

[18] These results are unchanged if delay measures calculated over the year following option introduction are examined instead.

firm specific news are reported in the following section. To capture the change in stock price adjustment efficiency related to option introduction, I contrast the various delay measures estimated over the year prior and following option introduction.

Panel A of Table V summarizes the change in the extended market model delay measures from the year prior to the year following option introduction, sorted by institutional ownership (proxy for short sale loan supply). Approximately 25% of the sample consisted of stocks with 0% institutional ownership (Nil INST GROUP). The remaining stocks were divided into three equal subsets (Low to High INST GROUPS) to form approximate institutional ownership quartiles. Moving incrementally between quartiles, the percentage of shares outstanding held by institutional investors increases by approximately 22%.

Focusing first on the extended market model, which considers improvement in market efficiency in relation to all market news, only stocks with low or nil institutional ownership (proxy for high short sale constraint levels) realize a significant improvement in stock price adjustment efficiency (negative ΔD). The extended neg market model, which examines the speed of stock price adjustment to negative market news, illustrates that the noted improvement in market efficiency occurs only for short sale constrained stocks, in response to negative new information. Collectively, these results support the conclusion that option introduction has a significant mitigating effect on short sale constraints. As a measure of the significance of this effect, prior to option introduction, short sale constrained stocks (Nil INST GROUP) adjust to negative information 19% more slowly than unconstrained stocks (High INST GROUP). Following option introduction that difference is reduced to 4%, indicating option introduction eliminates 79% of the price efficiency disparity between short sale constrained and unconstrained stocks (based on DM_{rsq}^{neg} values).

The results in Panel A of Table IV are subject to a series of limitations. First, although the results are robust to variation of the institutional ownership groupings, establishment of the breakpoints between groups can be criticized as arbitrary. Second, it is possible that institutional ownership proxies for a different stock characteristic, other than short sale loan supply, such as stock size, liquidity or investor inattention which is related to adjustment efficiency. It is noteworthy that the differential effect between positive and negative news unto itself significantly limits the range of alternative factors institutional ownership may be capturing. For example, it is difficult to imagine a circumstance where reduction of investor inattention or thin trading would result in an asymmetric improvement in efficiency for negative relative to positive news.

To more formally test the effect of option introduction on short sale constraints I complete cross-sectional regressions, regressing the change in stock price adjustment delay on institutional ownership (proxy for short sale loan supply) and control variables. To control for microstructure variables potentially also proxied by institutional ownership I include size, volatility and illiquidity. To control for investor inattention I also include analyst coverage, although size and illiquidity also serve as effective controls for investor inattention. Stocks potentially overlooked by investors tend to be small and illiquid with low turnover, analyst following and institutional investor interest.[19] As a final control I include the change in institutional ownership from one year prior to one year after option introduction as an increase in institutional investor interest could also be related to an improvement in stock price adjustment efficiency. The cross-sectional regression results are reported in Panel B of Table V, for each specification I report one model which includes the complete set of control variables and a second model which includes only control variables significant at conventional levels (α=0.05).

As short sale constraints are relaxed, new information is expected to be incorporated into stock prices more rapidly, resulting in a reduction in the delay measures (i.e. a negative ΔD). As short sale constraints are most binding for low institutional ownership stocks, a positive coefficient value is expected on INST. Further, as short sale constraints impede only the rate of incorporation of negative information in stock prices, a significant positive coefficient value is expected on INST in the DM Neg models while the INST coefficient in the DM Pos models is expected to be insignificant. I find results generally consistent with these expectations.

As I am primarily interested in the differential change in stock price adjustment delay in relation to negative relative to positive news I focus on the extended neg market model results which allow a cleaner interpretation. Focusing first on the $\Delta\, DM_{rsq}^{neg}$ measure, short sale constrained stocks experience a significant (t-stat 2.24) improvement in the speed of stock price adjustment to negative news

[19] Hou and Moskowitz (2005) also consider advertising spending and measures of remoteness (distance between the stock's headquarters and all U.S. airports for example) but these measures were not found to be significant determinants of stock price adjustment delay by them and thus are not considered.

following option introduction (positive coefficient value for INST variable). This relation is mirrored in the $\Delta\, DM_{se}^{neg}$ adjustment efficiency measure (t-stat 2.64). The relation between the change in the speed of stock price adjustment to positive information ($\Delta\, DM_{se}^{pos}$) and short sale constraints (proxied by institutional ownership) is insignificant at conventional levels (t-stat 1.03). [20] The selected control variables are generally insignificant with the exception of illiquidity. Stocks which realize post option improvement in stock price efficiency are characterized by high relative illiquidity in the year prior to option introduction. Overall, the results in Panel B of Table IV provide additional evidence of the significance of options to mitigate short sale constraints.

The short sale constraint proxy (institutional ownership) is a continuous variable in the market model regressions and the firm model regressions discussed further in the chapter. It would be expected that short sale constraints would bind below a certain share supply threshold and above that level, where supply likely exceeds demand, additional supply would be irrelevant. To test this expectation I replicate the market model analysis using an institutional ownership dummy in place of the continuous ownership variable, where the dummy variable is equal to one if institutional ownership is below a varying threshold and is otherwise equal to zero. I find that the institutional ownership dummy variable remains positive and significant where the ownership dummy variable is equal to one for firms with institutional ownership up to approximately 30% of shares outstanding. But, if

[20] The extended market model results are robust when replicated with the inclusion of 10 or 15 lags in the model.

stocks with no institutional holdings are excluded (0% institutional ownership) then the ownership dummy variable is insignificant across the majority of values less than 30%. These results support the expectation that a critical share loan supply threshold exists below which short sale constraints bind and that the critical value lies either at or close to 0% institutional ownership.

VI. Option Introduction and Stock Price Adjustment to Firm Specific News

The final analysis of the option introduction event study is to contrast post-option efficiency gains related to market-wide and firm specific news. The distinction between market and firm specific news is of interest as market-wide news consists predominantly of publicly available information where as firm specific news (proxied by excess stock returns) contains both private and public information. Miller (1977) argues that short sale constraints result in an irrational upward bias in prices due to trading by overly optimistic investors, thus negative public and private information would be equally excluded from prices. Conversely, Diamond and Verrecchia (1987) assume a rational framework in their model and argue that short sale constraints only impede the speed of stock price adjustment to negative private news. Thus, consistency in the change in stock price adjustment delay between market-wide news and firm specific news would be supportive of the Miller (1977) hypothesis. Alternatively, under the Diamond and Verrecchia (1987) hypothesis significant improvement in stock price adjustment should be isolated to the private information component of firm specific news.

Panel A of Table VI presents the univariate analysis of the extended firm model delay measures. The results are generally consistent with the univariate analysis of the extended market model with the exception of reduced significance in the extended neg firm model. For the Nil institutional ownership group the mean ΔDF_{rsq}^{neg} measure is marginally significant at conventional levels (t statistic value of 1.95 based on the null hypothesis that the mean = 0) and the mean ΔDF_{se}^{neg} is not statistically significant. When the mean delay measures of the extended firm models are compared to the corresponding extended market models, none of the delay measures between the two models are statistically unique (for both overall and neg models).[21]

These results are mirrored in the cross-sectional regression tests presented in Panel B of Table V in which the extended firm model delay measures are regressed on institutional ownership (proxy for short sale loan supply) and control variables. Recall, a negative ΔD indicates an improvement in the speed of stock price adjustment, thus a positive and significant INST coefficient would indicate an improvement in adjustment efficiency for short sale constrained stocks. For all models, with the exception of the ΔDF_{se}^{neg} delay measure model, the institutional ownership coefficient is positive and significant. Further, the institutional ownership coefficient in the ΔDF_{se}^{Pos} model is insignificant (t statistic 0.67) suggesting no significant post-option improvement in stock price adjustment

[21] As a robustness check, the correlation between market and excess stock returns was measured as a potential alternative source of commonality between the market and firm models is correlation between the two return series. The mean correlation coefficient across stocks between daily market and excess stock returns was 0.088, suggesting that correlation between the two return series is not a concern.

efficiency in relation to positive firm specific news. As noted for the extended market model, the control variables are generally insignificant with the exception of illiquidity and the change in institutional ownership. Stocks which realize a significant post option improvement in stock price adjustment tend to have high relative pre-option illiquidity and realize an increase in institutional ownership in the year following option introduction.

The collective results of the analysis of the effect of option introduction on short sale constraints can be summarized as follows. First, in relation to both firm specific and market-wide news, post-option improvement in the speed of stock price adjustment is isolated to low institutional ownership stocks (short sale constraint proxy) responding to negative news. No significant effect is noted in relation to the speed of adjustment to positive news for either firm specific or market-wide information. This result suggests options play a significant role in mitigating short sale constraints. Second, commonality in the improvement in price adjustment efficiency for both market-wide public and firm specific jointly private and public negative information suggests short sale constraint effects stem from an irrational optimism bias or another behavioral source as suggested by Miller (1977). Delayed incorporation of negative private information into stock prices likely also plays a role in short sale constraint effects but is not the exclusive source in a rational expectations framework, as suggested by Diamond and Verrecchia (1987). Unfortunately, within the model construct it is not possible to isolate the speed of adjustment to the private and public components of firm specific returns to explicitly test the relative contribution of each short sale constraint effect source to the overall effect.

VII. Robustness

Unlike stock exchanges, the decision to list an option is made at the discretion of the board of the option exchange and not at the discretion of the firm's board of directors. Assuming the primary motivation of an option exchange is to maximize the long term profitability of the exchange, the board of the option exchange will select stocks which are likely to generate the largest, long term trading volume (Mayhew and Mihov, 2004). Danielsen et al. (2007) examine stock characteristics which are predictive of option introduction and find that stocks with high market capitalization, improving liquidity and high abnormal volatility are favored for option listing. While it is unlikely option exchanges are attempting to select stocks with improving efficiency, it is possible they may indirectly do so based on their established selection criteria. This conclusion is supported by the increase in institutional ownership noted in Figure 1, which suggests stocks undergo an increase in investor recognition and /or popularity prior to option introduction.

To rule out the potential influence of endogenous stock characteristics related to option introduction not already considered as controls in the cross-sectional regressions, I create two control samples of non-optioned stocks matching the characteristics of the original sample (referred to forward as the test sample). To construct the first control sample I utilize the methods in Danielsen et al. (2007), who estimate a Cox Proportional Hazard Model on pooled monthly observations for all stocks classified as eligible for option listing but not yet optioned from 1993 to 2002. Danielsen et al. (2007) consider spread, abnormal spread, volatility, abnormal volatility, volume, abnormal volume, size and price as potential determinants of option introduction likelihood and determine that the spread, volatility and size variables are the most significant determinants.

To develop the pool of potential control stocks, from the universe of stocks tracked by the Center for Research in Security Prices (CRSP), I remove all stocks for which option trading was introduced from January 1973 – January 1997. Considering the remaining universe of CRSP stocks, each stock within the test sample is paired with a control stock matched on the following sorting variables from Danielsen et al. (2007): SIZE is calculated as market capitalization the day prior to option introduction, PCT SPREAD40-20 is the average daily closing percent spread from 40 to 20 trading days prior to option introduction, where closing percent spread is calculated as closing spread (ask - bid) divided by the midpoint between the closing bid and ask. ABSPREAD is calculated as PCT SPREAD40-20 divided by the average daily closing percent spread from 250 to 125 trading days prior to option introduction. ABSDRET and SDRET40-20 are the standard deviation of daily returns calculated in the same manner for the same timeframes. As it is desirable for the control sample to match the test sample in relation to short sale constraint levels, to the list of variable considered by Danielsen et al. (2007) I add INST, defined as the percentage of shares outstanding held by institutional investors in the option introduction quarter. All timeframes are in reference to the option introduction date of the test sample stock. The stock with the lowest equal weighted, mean percent difference of each of the six sorting variables was selected for the control sample, where percent difference was calculated as the absolute value of the difference between the control and test variable divided by the test variable. Once selected for the control sample, that stock was not eligible for future selection. The end result of the process is the construction of a control sample of 1145 stocks, each matched to a specific stock in the test sample based on the equal weighted percent difference in percent spread, abnormal percent spread, volatility, abnormal volatility, size and institutional ownership.[22] It is noteworthy that the selected sorting variables closely match the variables found to be significant determinants of stock price adjustment delay in Table III (size, volatility and illiquidity). Thus, the control sample is matched to the test sample based on both option introduction likelihood and determinants of price adjustment efficiency.

Panel A of Table VII presents the mean values of each of the six control methodology sorting variables for the test sample and control sample. The percent difference between the two samples is below 15% for each of the variables with the exception of institutional ownership (23.82%) and PCT SPREAD40-20 (30.34%). The difference in these two sorting variables suggests the control sample is more short sale constrained and has higher liquidity than the test sample. Differences between the control sample and test sample which bias towards a significant post-

[22] Bid and ask data were obtained from the Center for Research in Security Prices database. Bid and ask data was not available for the entire sample, reducing the number of stocks available for matching in the control sample to 1145 (a reduction of 34%). The general proportionality of stocks in each of the institutional ownership quartiles was not significantly altered, for example the number of stocks in the Nil institutional ownership group was reduced by 33% and the other groups experienced similar reductions. Thus, the proportionality of short sale constraint levels in the reduced sample is considered equivalent to the original.

option change in stock price adjustment (lower size and institutional ownership and higher volatility) are more common than differences in variables which suggest the opposite (higher liquidity). Thus, if anything, the differences between the control sample and test sample bias against the finds of this paper. Additionally, Danielsen et al. (2007) note the dynamics of spread as the single most important determinant of option introduction likelihood and on this variable the difference between the control and test sample is less than 1%. As an additional robustness check I constructed a second control sample matched to the test sample based only on size and institutional ownership. Using the second methodology I am able to match the entire sample (1732 stocks) and the percent difference of SIZE and INST is reduced to 0.8% and 21.3% respectively. As the results are consistent between the two control samples, only results related to the first control sample are presented.

Panel B of Table VI presents the univariate analysis of the change in delay measures for the control sample. The institutional ownership groups (INST GROUP) are established in the same manner as the test sample, allocating stocks with zero percent institutional ownership to the first group and then equally allocating the remaining stocks to the three remaining groups to establish approximate quartiles. Recall, a reduction in each delay measure reflects an improvement in stock price adjustment efficiency. Over the timeframe of option introduction, for each of the eight delay measures, no significant change in stock price adjustment efficiency is noted for stocks with high short sale constraint levels (Nil and Low INST GROUP). Stocks with low short sale constraint levels (Med and High INST GROUP) realize a significant decrease in stock price adjustment efficiency for the majority of the price delay measures, with the exception of the $\Delta\, DF_{rsq}^{neg}$ measure for the Med INST GROUP which reflects an improvement in adjustment efficiency. The decline in efficiency noted for the Med and High INST groups in the control sample is also noted to a lesser extent in the test sample. This trend is consistent with a peak and associated decline in investor interest associated with market timing of option introductions or stocks with similar characteristics commonly associated with listing. As discussed in the next paragraph, the results related to the $\Delta\, DF_{rsq}^{neg}$ measure for the Med INST GROUP are insignificant in the multivariate analysis with the inclusion of control variables. Thus, the trend can be associated with other stock characteristics than short sale constraints.

Panel C of Table VI reports the results of the multivariate analysis of the control sample, in which the adjustment delay measures are regressed on institutional ownership and a series of control variables. After controlling for size, volatility, illiquidity and analyst coverage, no significant relation is noted between the change in stock price adjustment efficiency and short sale constraint levels (proxied by institutional ownership) in relation to both positive and negative information.[23] Collectively, these results reflect no significant reduction in short sale constraint levels for the control sample over the timeframe of option introduction of the test sample.

[23] As the extended negative and positive delay measures provide the cleanest test of the change in stock price adjustment in relation to negative news only results for those variables are reported. Results in relation to the extend market model and extended firm model are consistent with the results presented in Panel B of Table VI and are available upon request.

As a final robustness check I replicate the market and firm models to check for potential model biases related to the OLS regression's assumption of recursivity. An issue in the cross-sectional regressions is that institutional ownership at the time of option introduction may be affected by the other independent variables included in the regression. I adopt a three-stage model to correct for the potential endogeneity of institutional ownership. In the first stage, I model institutional ownership at the time of option introduction as a function of pre-introduction firm characteristics, as shown in equation 2.15 (all variables are as previously described in Table III).

$$INST_{i,t} = \alpha_i + SIZE_{i,t-1} + TURN_{i,t-1} + VOL_{i,t-1} + ILLIQ_{i,t-1} + ANAL_{i,t-1} + EMPLOY_{i,t-1} + \varepsilon_i \qquad (2.15)$$

In the second stage, I model the change in institutional ownership from the year before to the year after option introduction as a function of ownership in the listing quarter and average daily turnover of the stock over the year prior to option listing, as shown in equation 2.16.[24]

$$\Delta INST_{i,t} = \alpha_i + INST_{i,t-1} + TURN_{i,t-1} + \varepsilon_i \qquad (2.16)$$

In the third stage I replicate the firm and market models, replacing the institutional ownership variables with the residuals from the models in equation 2.15 and 2.16. Using the orthoganalized proxies for institutional ownership and the

[24] All of the variables selected for inclusion in the first and second stages are significant using an alpha value of 0.05.

change in institutional ownership I find the same results, specifically the residual institutional ownership coefficient is positive and statistically significant for both the firm and market negative extended models and insignificant for the positive extended models. In the interest of brevity these results are not reported but are available upon request.

VIII. Options and the 2008 Short Sale Ban

Between September 19 and October 6, 2008 the SEC banned short sale transactions for 986 financial sector stocks in an effort to stabilize the market during the 2008 global financial crisis. The ban was lifted after 13 trading days on October 8t, 2008. Despite the ban, on average S&P 500 Index listed financial sector stocks lost 24% of value during the ban period. In order to enable continued function of option markets, option bookmakers were exempt from the short sale ban and thus were able to continue to hedge their put option exposure via short sales. Given the continued ability of investors to synthetically short a subset of financial sector stocks during the short sale ban, the ban provides a natural experiment to test the potential for options to mitigate legislatively imposed short sale constraints.

Given the short duration of the short sale ban it is not feasible to conduct the stock price adjustment delay analysis utilized to examine market efficiency surrounding option introduction. Instead, I examine the cumulative returns to optioned and non-optioned stocks during the banned period. If options act to mitigate short sale constraints it would be expected that negative information would be incorporate more freely into optioned stocks resulting in relative overpricing of non-optioned stocks. Figure 2 displays cumulative abnormal returns to optioned and non-optioned stocks commencing two weeks before and concluding two weeks after the ban period. *CUMRET* is calculated as the compounded value of the daily stock return commencing September 2, 2008 (day 0),

$$CUMRET_i = \prod_{t=-0}^{33} (1 + R_{i,t}) \qquad (2.17)$$

where $R_{i,t}$ is the return for stock *i* on date *t*.

In the two weeks before and after the ban period, optioned and non-optioned stocks exhibit highly similar return patterns. Focusing on the actual short sale ban period, calculating cumulative returns from September 22 to October 8, 2008 optioned and non-optioned stock prices dropped 29% and 16% in value, respectively.[25] Although the 13% difference in cumulative returns is statistically significant, it is possible that aggregate differences in stock characteristics or variations in the sensitivity to financial sector news between the two sub-samples could also explain this difference. Table VIII presents cross-sectional regression results, regressing the cumulative return on October 8, 2008 of each stock on an option dummy and control variables.[26] As controls I include market capitalization the day prior to the ban, turnover and illiquidity calculated over the year prior to the

[25] I commence calculation of cumulative returns on September 22, 2008 as at that time over 90% of the stocks on the final short sale ban list had been banned. Results are similar if cumulative returns are calculated starting on September 19, 2008 when the initial ban list was released.

[26] The option dummy equals one if the stock is listed on the CBOE, Philadelphia Stock and Option Exchange or the NYSE / AMEX Option Exchange and is otherwise equal to zero. The option listings were obtained from the website of each option exchange.

short sale ban and the book to market ratio calculated as book value on the day of the firm's 2007 fiscal year end divided by market capitalization.[27] Although all of the stocks in the ban sample were featured prominently in the financial press at the time of the ban, I also include institutional holdings as of August 30, 2008 as an additional control for investor inattention.[28] To control for variations in stock sensitivity to financial sector news I include the beta coefficient from equation 2.18 calculated over the year prior to the short sale ban:

$$r_{i,t} - rf_t = \alpha_i + \beta_{BAN,i}(R_{BAN,t} - rf_t) + \varepsilon_i \tag{2.18}$$

where $r_{i,t}$ is the return to stock i on day t, rf is the risk free rate on day t and $R_{BAN,t}$ is the equal weighted average return to the short sale ban sample also on day t.[29]

Reviewing the regression output in Table VIII, the option dummy coefficient is negative and significant at conventional levels (t-statistic 5.44) indicating optioned stocks realized lower cumulative returns during the short sale ban. The majority of the control variables are insignificant, but the institutional ownership variable is positive and marginally significant (t-statistic 1.53) suggesting stocks with higher institutional ownership realized higher returns. The banned stock index beta (β_{BAN}) is negative and significant (t-statistic 11.58) indicating, as intuitively would be

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27 Illiquidity and turnover are calculated in the same manner as in the option introduction tests. As in the option introduction analysis, the average book to market value of the sample was used for stocks missing book value data on Compustat.

28 Institutional ownership data were obtained from the Short Squeeze database.

29 Daily risk free rate data were obtained from Kenneth French's website.

expected, that stocks with high financial sector sensitivity realized significantly lower cumulative returns during the short sale ban. Collectively the results of the short sale ban analysis suggest negative information was incorporated more free into optioned stocks during the short sale ban as options acted as an effective substitute to short sales, allowing investors to realize a synthetic short position.

IX. Concluding Remarks

In this paper I re-examine the potential for options to mitigate short sale constraints, conducting two event studies focusing on option introductions and the short sale ban in September, 2008. By examining cross-sectional variation in post-option efficiency improvements, across short sale constraint levels, this paper documents a significant improvement in market efficiency for a subsection of stocks with low short sale loan supply. By mitigating short sale constraints, option introduction is found to eliminate 79% of the stock price adjustment efficiency disparity between short sale constrained and unconstrained stocks in relation to negative information. I also document that during the short sale ban negative information was incorporated into optioned stock prices more freely, providing further evidence that options act as a substitute for short sales.

While the effect of short sale constraints on market efficiency is generally well documented in the finance literature, the potential for options to mitigate short sale constraints is less well understood and disputed. Beyond contributing to this ongoing debate, the findings of this paper have significant policy implications. This paper does not examine the overall potential effects of the short sale ban, leaving that analysis to other research. But, by exempting option bookmakers from the short sale ban, any potential effect of the ban was likely diminished by option trading providing an alternative mechanism for investors to realize a synthetic short position. When considering future restrictive action in relation to short sales, policy makers should consider the joint effects of short sales and option markets and the potential for them to act as compliments when both are functioning and as substitutes in the absence of each other.

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Table I

Summary statistics for the option introduction sample

Price is the stock price and Shares is the number of shares outstanding both on the day before option introduction. Volatility is the average standard deviation of daily stock price returns over the year prior to option introduction. Turnover is the average daily turnover in the year preceding option introduction calculated as daily trading volume divided by shares outstanding. Size is market capitalization on the day prior to option introduction. Book is book value per share reported for the year which the underlying stock was listed on the option exchange. BK/MK is book value of the stock in the year of option introduction divided by market capitalization of the stock on the day prior to option introduction.

	Price ($/share)	Shares (million)	Volatility	Turnover (%)	Size (million $)	Book ($/share)	BK/MK
N	1732	1732	1732	1732	1732	1719	1719
Mean	26.45	26.74	0.0294	0.70	805	17.23	1.01
Median	23.25	17.07	0.0279	0.51	470	4.31	0.19
Std Dev	16.65	30.89	0.0119	0.61	103	153.66	11.64
Q3	33.50	29.78	0.0369	0.95	937	8.80	0.41
Q1	15.60	11.22	0.0202	0.28	257	1.80	0.08

Table II

Short sale ban sample summary statistics

Price is the stock price and Shares is the number of shares outstanding in millions both on September 18, 2008, the day before the commencement of the short sale ban. Volatility and turnover are calculated over the year prior to the short sale ban, where volatility is the average standard deviation of daily stock price returns and turnover is the average daily ratio of shares traded divided by shares outstanding. Size is market capitalization on the day prior to the short sale ban. Book is book value per share reported for the firm's fiscal year end in 2007. BK/MK is book value divided by market capitalization.

Panel A: Descriptive statistics of short sale banned stocks with no option listing

	Price ($/share)	Shares (million)	Volatility	Turnover (%)	Size (million $)	Book ($/share)	BK/MK
N	533	533	533	533	533	456	456
Mean	266.37	19.20	0.034	1.30	768	16.15	1.55
Median	10.50	7.67	0.032	0.11	768	11.52	1.12
Std Dev	5546.76	111.06	0.014	17.77	6795	30.98	1.66
Q3	17.37	13.75	0.041	0.28	206	15.98	1.71
Q1	7.03	4.10	0.026	0.053	37	8.16	0.78

Panel B: Descriptive statistics of short sale banned stocks listed on an option exchange

	Price ($/share)	Shares (million)	Volatility	Turnover (%)	Size (million $)	Book ($/share)	BK/MK
N	383	383	383	383	383	339	339
Mean	30.29	390	0.037	2.70	10,171	20.69	1.96
Median	23.90	71	0.033	0.96	1,918	16.12	0.87
Std Dev	30.33	107	0.019	12.47	24,968	16.81	3.87
Q3	38.14	268	0.042	1.60	6,525	26.09	1.56
Q1	13.19	35	0.025	0.66	624	10.18	0.53

Table III

Determinant variable correlation matrix

INST is the percentage of total shares outstanding held by institutional investors in the quarter of option introduction. SIZE is market capitalization the day prior to option introduction. BK/MK is book value of the stock in the year of option introduction divided by market capitalization of the stock on the day prior of option introduction. TURN is the average daily turnover in the year preceding option introduction calculated as daily trading volume divided by shares outstanding. VOL is the average standard deviation of daily stock price returns over the year prior to option introduction. ILLIQ is the average weekly Amihud Illiquidity Ratio (Amihud, 2002) over the year preceding option introduction calculated as the absolute weekly return divided by the dollar value of weekly trading volume. ANAL is the number of analysts which cover the stock in the year of option introduction and EMPLOY is the number of people employed by the firm in the fiscal year end of option introduction. Correlation values which appear in bold face are significant at conventional levels (α= 0.05).

	INST	BK/MK	TURN	ILLIQ	VOL	SIZE	ANAL	EMPLOY
INST	1							
BK/MK	-0.0169	1						
TURN	0.0131	-0.00635	1					
ILLIQ	**-0.112**	0.0179	-0.0033	1				
VOL	**-0.123**	0.0442	**0.560**	**0.263**	1			
SIZE	-0.0450	-0.0292	**-0.341**	**-0.133**	**-0.476**	1		
ANAL	**0.229**	-0.00684	**-0.0729**	**-0.0699**	**-0.133**	**0.077 8**	1	
EMPLOY	**-0.049**	-0.00332	**-0.0681**	**-0.0529**	**-0.206**	**0.281**	0.0176	1

Table IV

Determinants of stock price adjustment delay

Table IV reports cross sectional regression results of stock price adjustment delay measures regressed on determinant variables. The table reports standardized coefficient values for each variable with t-statistic values reported below. Coefficients which are significant at conventional levels (α=0.05) are reported in bold face. Delay measures are calculated over the year prior to option introduction. The notation DM denotes delay measures that are determined using lagged market returns and DF denotes delay measures that are determined using lagged excess stock returns. INST is the percentage of total shares outstanding held by institutional investors in the quarter of option introduction. SIZE is market capitalization the day prior to option introduction. BK/MK is book value of the stock in the year of option introduction divided by market capitalization of the stock on the day prior to option introduction. TURN is average daily turnover in the year preceding option introduction calculated as daily trading volume divided by shares outstanding. VOL is the average standard deviation of stock price returns over the year prior to option introduction. ILLIQ is the average weekly Amihud Illiquidity Ratio (Amihud, 2002) for the year preceding option introduction calculated as the absolute weekly return divided by the dollar value of weekly trading volume. ANAL is the number of analysts which cover the stock in the year of option introduction and EMPLOY is the number of people employed by the firm in the fiscal year end of option introduction.

Dependent Variable	Intercept	INST	SIZE	BK/MK	TURN	VOL	ILLIQ	ANAL	EMPLOY	adj R^2
DM_{rsq} *Pre*	**0.00**	**-0.10**	-0.023	-0.001	-0.019	0.064	**0.16**	-0.038	-0.029	0.0542
	(14.29)	(4.03)	(0.83)	(0.05)	(0.66)	(1.99)	(6.49)	(1.58)	(1.16)	
DM_{rsq} *Pre*	**0.00**	**-0.10**				**0.074**	**0.16**			0.0539
	(17.23)	(4.39)				(3.03)	(6.78)			
DM_{se} *Pre*	**0.00**	**-0.061**	**-0.13**	-0.001	-0.042	0.026	**0.13**	**-0.062**	**-0.056**	0.0583
	(33.77)	(2.51)	(4.87)	(0.04)	(1.45)	(0.79)	(5.19)	(2.57)	(2.26)	
DM_{se} *Pre*	**0.00**	**-0.064**	**-0.13**				**0.14**	**-0.061**	**-0.059**	0.0588
	(70.77)	(2.67)	(5.28)				(5.70)	(2.55)	(2.42)	
DM_{rsq} Neg Pre	**0.00**	**-0.091**	**-0.14**	-0.021	-0.057	0.054	**0.13**	**-0.017**	**-0.042**	0.0583
	(16.16)	(3.73)	(5.10)	(0.92)	(1.97)	(1.68)	(5.29)	(0.71)	(1.72)	
DM_{rsq} Neg Pre	**0.00**	**-0.10**	**-0.13**				**0.14**			0.0547
	(35.59)	(4.36)	(5.69)				(6.04)			
DM_{se} Neg Pre	**0.00**	**-0.072**	**-0.11**	-0.006	-0.029	0.063	0.035	-0.015	0.028	0.0245
	(35.70)	(2.88)	(4.00)	(0.26)	(0.98)	(1.93)	(1.39)	(0.60)	(1.14)	
DM_{se} Neg Pre	**0.00**	**-0.081**	**-0.10**			0.55				0.0245
	(36.27)	(3.37)	(3.84)			(1.99)				

Table IV is continued on the next page.

Table IV

Determinants of stock price adjustment delay (continued from previous page)

Dependent Variable	Intercept	INST	SIZE	BK/MK	TURN	VOL	ILLIQ	ANAL	EMPLOY	adj R^2
DF_{rsq} Pre	**0.00**	**-0.091**	**-0.11**	-0.010	-0.036	**0.093**	**0.20**	-0.041	-0.0023	0.0959
	(13.57)	(3.78)	(3.94)	(0.42)	(1.28)	(2.95)	(8.08)	(1.72)	(0.99)	
DF_{rsq} Pre	**0.00**	**-0.10**	**-0.11**			**0.078**	**0.20**			0.0949
	(13.58)	(4.31)	(4.20)			(2.90)	(8.49)			
DF_{se} Pre	**0.00**	**-0.071**	**-0.13**	-0.032	-0.009	**0.088**	**0.092**	**-0.055**	-0.026	0.0638
	(33.94)	(2.91)	(4.67)	(1.40)	(0.32)	(2.73)	(3.75)	(2.28)	(1.04)	
DF_{se} Pre	**0.00**	**-0.070**	**-0.13**			**0.084**	**0.094**	**-0.055**		0.0637
	(34.17)	(2.87)	(4.98)			(3.07)	(3.88)	(2.28)		
DF_{rsq} Neg Pre	**0.00**	**-0.10**	**-0.12**	-0.01	-0.042	**0.13**	**0.078**	-0.010	0.046	0.0627
	(14.50)	(4.38)	(4.37)	(0.37)	(1.46)	(4.13)	(3.16)	(0.43)	(1.88)	
DF_{rsq} Neg Pre	**0.00**	**-0.11**	**-0.11**			**0.11**	**0.085**			0.0619
	(14.97)	(4.81)	(3.97)			(3.83)	(3.50)			
DF_{se} Neg Pre	**0.00**	-0.015	-0.052	-0.019	-0.018	0.012	0.002	**-0.072**	-0.009	0.0058
	(37.38)	(0.58)	(1.85)	(0.81)	(0.59)	(0.35)	(0.09)	(2.91)	(0.37)	
DF_{se} Neg Pre	**0.00**		**-0.054**					**-0.076**		0.0082
	(113.84)		(2.26)					(3.17)		

Table V

Option introduction and the speed of stock price adjustment to market news

Panel A: Extended market model mean delay measures

Panel A presents the average change in the extended market model stock price adjustment delay measures following option introduction sorted by institutional ownership. Institutional ownership (INST) is the percentage of total shares outstanding held by institutional investors in the quarter of option introduction. Mean ΔD_{rsq} and mean ΔD_{se} are the mean changes in each delay measure between the year before and the year after option introduction for each subset.

			Extended Market Model		***Extended Neg Market Model***	
INST GROUP	N	Mean INST	Mean ΔDM_{rsq}	Mean ΔDM_{se}	Mean ΔDM_{rsq}^{neg}	Mean ΔDM_{se}^{neg}
Nil	442	0.00	-0.0248**	-0.0264**	-0.0209**	-0.1022**
Low	430	0.22	-0.0247**	-0.0158**	-0.0231**	0.0050
Med	430	0.45	-0.0021	0.04060	-0.00021	-0.00066
High	430	0.66	0.0100	0.01712	0.0055	0.0066

** Change in mean significant α= 0.05
* Change in mean significant α= 0.10

Panel B: Extended market model regression results

Panel B reports cross-sectional regression results of the change in stock price adjustment delay measures regressed on institutional ownership (proxy for short sale loan supply) and control variables. The change in each delay measure is calculated as the difference in each measure over the year before and after option introduction. The table reports standardized coefficient values for each variable with t-statistic values reported below. Coefficients which are significant at conventional levels (α=0.05) are reported in bold face. INST is the percentage of total shares outstanding held by institutional investors in the quarter of option introduction. ΔINST is the change in institutional ownership one year prior to one year following option introduction. SIZE is market capitalization the day prior to option introduction. BK/MK is book value of the stock in the year of option introduction divided by market capitalization of the stock on the day prior to option introduction. TURN is average daily turnover in the year preceding option introduction calculated as daily trading volume divided by shares outstanding. VOL is the average standard deviation of stock price returns over the year prior to option introduction. ILLIQ is the average weekly Amihud Illiquidity Ratio (Amihud, 2002) for the year preceding option introduction calculated as the absolute weekly return divided by the dollar value of weekly trading volume. ANAL is the number of analysts which cover the stock in the year of option introduction and EMPLOY is the number of people employed by the firm in the fiscal year end of option introduction.

Dependent Variable	Intercept	INST	ΔINST	SIZE	BK/MK	TURN	VOL	ILLIQ	ANAL	EMPLOY	adj R^2
Δ DM_{rsq}	0.00	0.046	-0.044	-0.039	0.021	0.035	-0.039	**-0.087**	-0.011	0.007	0.109
	(0.37)	(1.79)	(1.821)	(1.38)	(0.88)	(1.19)	(1.17)	(3.45)	(0.46)	(0.27)	
Δ DM_{rsq}	0.00	0.043						**-0.092**			0.100
	(1.78)	(1.77)						(3.81)			
Δ DM_{se}	0.00	0.017	-0.025	0.032	0.009	0.040	-0.021	**-0.061**	0.013	0.011	0.0023
	(0.58)	(0.65)	(1.01)	(1.12)	(0.39)	(1.33)	(0.64)	(2.41)	(0.51)	(0.44)	
Δ DM_{se}	**0.00**	0.019						**-0.064**			0.0036
	(2.31)	(0.80)						(2.65)			

Panel B of Table V is continued on the next page.

Panel B: Extended market model regression results (continued from previous page)

Dependent Variable	Intercept	INST	ΔINST	SIZE	BK/MK	TURN	VOL	ILLIQ	ANAL	EMPLOY	adj R^2
Δ DM_{rsq} Neg	0.00	**0.060**	-0.034	0.042	0.058	0.039	0.010	**-0.073**	-0.012	0.075	0.0076
	(1.32)	(2.36)	(1.38)	(1.50)	(2.41)	(1.32)	(0.30)	(2.90)	(0.48)	(3.01)	
Δ DM_{rsq} Neg	**-0.00**	**0.051**			**0.055**			**-0.081**	0.063		0.014
	(2.03)	(2.13)			(2.31)			(3.38)	(2.63)		
Δ DM_{se} Neg	0.00	**0.053**	0.026	0.022	**0.049**	0.002	-0.008	-0.003	-0.016	-0.027	0.0019
	(1.04)	(2.07)	(1.06)	(0.78)	(2.04)	(0.07)	(0.25)	(0.12)	(0.66)	(1.06)	
Δ DM_{se} Neg	**0.00**	**0.055**			**0.049**						0.0041
	(2.77)	(2.28)			(2.04)						
Δ DM_{se} Pos	**-0.00**	0.027	-0.009	**0.060**	0.018	0.025	0.011	-0.002	0.010	-0.013	-0.001
	(2.27)	(1.06)	(0.37)	(2.11)	(0.74)	(0.84)	(0.33)	(0.09)	(0.41)	(0.52)	
Δ DM_{se} Pos	**0.00**	0.025									0.00
	(2.98)	(1.03)									

Table VI

Option introduction and the speed of stock price adjustment to firm specific news

Panel A: Extended firm model mean delay measures

Panel A presents the average change in the extended market model stock price adjustment delay measures following option introduction sorted by institutional ownership. Institutional ownership (INST) is the percentage of total shares outstanding held by institutional investors in the quarter of option introduction. Mean ΔD_{rsq} and mean ΔD_{se} are the mean changes in each delay measure between the year before and the year after option introduction for each subset.

			Extended Firm Model		***Extended Neg Firm Model***	
INST GROUP	N	Mean INST	Mean ΔDF_{rsq}	Mean ΔDF_{se}	Mean ΔDF_{rsq}^{neg}	Mean ΔDF_{se}^{neg}
Nil	442	0.00	-0.0291**	-0.0557**	-0.0181*	-0.0024
Low	430	0.22	-0.0301**	-0.0355	-0.0163**	-0.0231
Med	430	0.45	-0.0001	0.0240	0.00358	0.00302
High	430	0.66	0.0095	0.0272	0.0131	-0.0090

** Change in mean significant α= 0.05
* Change in mean significant α= 0.10

Panel B: Extended firm model regression results

Panel B reports cross-sectional regression results of the change in stock price adjustment delay measures regressed on institutional ownership (proxy for short sale loan supply) and control variables. The change in each delay measure is calculated as the difference in each measure over the year before and after option introduction. The table reports standardized coefficient values for each variable with t-statistic values reported below. Coefficients which are significant at conventional levels (α=0.05) are reported in bold face. INST is the percentage of total shares outstanding held by institutional investors in the quarter of option introduction. Δ INST is the change in institutional ownership one year prior to one year following option introduction. SIZE is market capitalization the day prior to option introduction. BK/MK is book value of the stock in the year of option introduction divided by market capitalization of the stock on the day prior to option introduction. TURN is average daily turnover in the year preceding option introduction calculated as daily trading volume divided by shares outstanding. VOL is the average standard deviation of stock price returns over the year prior to option introduction. ILLIQ is the average weekly Amihud Illiquidity Ratio (Amihud, 2002) for the year preceding option introduction calculated as the absolute weekly return divided by the dollar value of weekly trading volume. ANAL is the number of analysts which cover the stock in the year of option introduction and EMPLOY is the number of people employed by the firm in the fiscal year end of option introduction.

Dependent Variable	Intercept	INST	ΔINST	SIZE	BK/MK	TURN	VOL	ILLIQ	ANALY	EMPLOY	adj R^2
Δ DF_{rsq}	0.00	**0.053**	**-0.057**	-0.052	0.026	0.008	-0.040	**-0.10**	0.009	0.009	0.0162
	(0.65)	(2.11)	(2.33)	(1.85)	(1.07)	(0.29)	(1.21)	(4.01)	(0.37)	(0.37)	
Δ DF_{rsq}	0.00	**0.062**	**-0.056**					**-0.10**			0.0167
	(1.71)	(2.53)	(2.33)					(4.36)			
Δ DF_{se}	0.00	0.042	-0.034	-0.039	0.045	0.014	-0.031	**-0.079**	0.030	0.005	0.0102
	(0.43)	(1.65)	(1.41)	(1.39)	(1.88)	(0.49)	(0.93)	(3.12)	(1.25)	(0.20)	
Δ DF_{se}	0.00	**0.048**						**-0.083**			0.0090
	(1.36)	(1.99)						(3.46)			
Δ DF_{rsq} Neg	0.00	**0.070**	**-0.090**	-0.008	0.043	0.049	-0.060	**-0.050**	-0.026	-0.041	0.0137
	(0.24)	(2.77)	(3.73)	(0.29)	(1.82)	(1.67)	(1.83)	(1.99)	(1.07)	(1.63)	
Δ DF_{rsq} Neg	**0.00**	**0.071**	**-0.088**					**-0.062**			0.0142
	(2.03)	(2.94)	(3.65)					(2.60)			

Panel B: Extended firm model regression results (continued from previous page)

Dependent Variable	Intercept	INST	ΔINST	SIZE	BK/MK	TURN	VOL	ILLIQ	ANALY	EMPLOY	adj R^2
Δ DF_{se} Neg	0.00 (0.02)	-0.018 (0.71)	-0.010 (0.42)	-0.015 (0.51)	0.048 (1.98)	0.002 (0.06)	-0.002 (0.07)	0.013 (0.52)	0.048 (1.94)	0.009 (0.34)	-0.0002
Δ DF_{se} Neg	0.00 (0.17)	-0.010 (0.43)									-0.0005
Δ DF_{se} Pos	0.00 (1.95)	0.014 (0.54)	-0.011 (0.43)	0.017 (0.60)	-0.001 (0.05)	-0.011 (0.37)	0.022 (0.67)	-0.022 (0.87)	0.032 (1.28)	0.043 (1.71)	-0.0009
Δ DF_{se} Pos	**0.00** (3.12)	0.002 (0.67)									-0.0003

Table VII

Control sample stock price adjustment delay analysis

Panel A: Comparison of mean sorting variables

Panel A presents the average value of the sorting variables used to establish the control sample for the test and control samples. SIZE is market capitalization the day before option introduction. INST is the percentage of total shares outstanding held by institutional investors in the quarter of option introduction. PCT SPREAD40-20 is the average daily closing percent spread from 40 to 20 trading days prior to option introduction, where closing percent spread is calculated as closing spread (ask - bid) divided by the midpoint between the closing bid and ask. ABSPREAD is calculated as PCT SPREAD40-20 divided by the average daily closing spread from 250 to 125 trading days prior to option introduction. ABSDRET and SDRET40-20 are the standard deviation of daily returns calculated in the same manner for the same timeframes. All timeframes are in reference to the option introduction date of the test sample. % Diff is calculated as the absolute difference of the control and test sample divided by the test sample.

Sample	SIZE	INST	ABSPREAD	PCT SPREAD40-20	ABSDRET	SDRET40-20
Control	483628.4	0.254	0.859	0.025	0.958	0.027
Test	562864.7	0.334	0.861	0.019	1.054	0.032
% Diff	14.08%	23.82%	0.15%	30.34%	9.02%	14.48%

Panel B: Mean delay measures

Panel B presents the average change in the stock price adjustment delay measures for the control sample sorted by institutional ownership. Institutional ownership (INST) is the percentage of total shares outstanding held by institutional investors in the quarter of option introduction. Mean ΔD_{rsq} and mean ΔD_{se} are the mean changes in each delay measure between the year before and the year after option introduction of the test sample for each subset.

			Extended Market Model				*Extended Firm Model*			
INST GROUP	N	Mean INST	Mean ΔDM_{rsq}	Mean ΔDM_{se}	Mean ΔDM_{rsq}^{neg}	Mean ΔDM_{se}^{neg}	Mean ΔDF_{rsq}	Mean ΔDF_{se}	Mean ΔDF_{rsq}^{neg}	Mean ΔDF_{se}^{neg}
Nil	239	0.00	-0.0089	-0.0097	0.0028	0.0120	0.010	0.0010	-0.015	0.032
Low	302	0.086	0.0095	0.0074	0.022	0.013	0.016	0.027	-0.019	-0.020
Med	302	0.31	0.068**	0.118**	0.018	0.0034	0.052**	0.059*	-0.045**	0.043*
High	302	0.57	0.056**	0.147**	0.023*	-0.0066	0.063**	0.088**	0.0074	0.018

** Change in mean significant α= 0.05
* Change in mean significant α= 0.10

Panel C: Control sample regression results

Panel C reports cross sectional regression results of the change in stock price adjustment delay measures regressed on institutional ownership (proxy for short sale loan supply) and control variables. The change in each delay measure is calculated as the difference in each measure over the year before and after option introduction. The table reports standardized coefficient values for each variable with t-statistic values reported below. Coefficients which are significant at conventional levels (α=0.05) are reported in bold face. INST is the percentage of total shares outstanding held by institutional investors in the quarter of option introduction. ΔINST is the change in institutional ownership one year prior to one year following option introduction. SIZE is market capitalization the day prior to option introduction. BK/MK is book value of the stock in the year of option introduction divided by market capitalization of the stock on the day prior to option introduction. TURN is average daily turnover in the year preceding option introduction calculated as daily trading volume divided by shares outstanding. VOL is the average standard deviation of stock price returns over the year prior to option introduction. ILLIQ is the average weekly Amihud Illiquidity Ratio (Amihud, 2002) for the year preceding option introduction calculated as the absolute weekly return divided by the dollar value of weekly trading volume. ANAL is the number of analysts which cover the stock in the year of option introduction and EMPLOY is the number of people employed by the firm in the fiscal year end of option introduction.

Dependent Variable	Intercept	INST	ΔINST	SIZE	BK/MK	TURN	VOL	ILLIQ	ANAL	EMPLOY	adj R^2
Δ DM_{rsq} Neg	**0.00**	-0.006	0.022	-0.052	0.018	-0.019	**-0.086**	-0.017	0.023	-0.015	0.0051
	(2.68)	(0.17)	(0.72)	(1.75)	(0.62)	(0.64)	(2.74)	(0.56)	(0.68)	(0.49)	
Δ DM_{rsq} Neg	**0.00**	0.011					**-0.089**				0.0065
	(2.26)	(0.39)					(3.00)				
Δ DM_{se} Neg	0.00	-0.003	-0.053	-0.030	-0.023	-0.006	-0.040	-0.013	-0.016	-0.013	-0.0011
	(1.70)	(0.08)	(1.78)	(1.03)	(0.77)	(0.19)	(1.27)	(0.42)	(0.47)	(0.45)	
Δ DM_{se} Neg	0.00	-0.011									-0.0008
	(0.63)	(0.37)									
Δ DM_{se} Pos	**0.00**	0.005	-0.006	**-0.089**	-0.007	0.006	-0.041	0.011	-0.058	0.009	0.0038
	(2.21)	(0.14)	(0.21)	(2.98)	(0.22)	(0.21)	(1.32)	(0.34)	(1.69)	(0.30)	
Δ DM_{se} Pos	**0.00**	-0.021		**-0.086**							0.0060
	(2.01)	(0.71)		(2.91)							

Panel C of Table VII is continued on the next page.

Panel C: Control sample regression results (continued from previous page)

Dependent Variable	Intercept	INST	ΔINST	SIZE	BK/MK	TURN	VOL	ILLIQ	ANAL	EMPLOY	adj R^2
Δ DF$_{rsq}$ Neg	**0.00**	0.033	-0.019	-0.027	0.012	0.023	**-0.082**	-0.022	0.060	0.036	0.0065
	(3.40)	(0.95)	(0.63)	(0.90)	(0.40)	(0.77)	(2.61)	(0.71)	(1.76)	(1.21)	
Δ DF$_{rsq}$ Neg	**0.00**	0.025					**0.078**		0.002		0.0085
	(4.20)	(0.76)					(2.64)		(1.93)		
Δ DF$_{se}$ Neg	**0.00**	0.021	-0.018	0.005	-0.005	-0.058	**-0.096**	0.009	-0.017	-0.045	0.0079
	(2.29)	(0.61)	(0.62)	(0.18)	(0.17)	(1.93)	(3.06)	(0.29)	(0.50)	(1.48)	
Δ DF$_{se}$ Neg	**0.00**	0.013				**-0.061**	**-0.093**				0.0105
	(2.34)	(0.51)				(2.07)	(3.18)				
Δ DF$_{se}$ Pos	**0.00**	0.018	-0.017	**-0.075**	0.044	-0.015	-0.060	-0.047	-0.041	-0.035	0.0077
	(2.80)	(0.54)	(0.58)	(2.49)	(0.15)	(0.52)	(1.91)	(1.49)	(1.22)	(1.17)	
Δ DF$_{se}$ Pos	**0.00**	0.0003		**-0.077**			**-0.074**				0.0078
	(2.52)	(0.01)		(2.61)			(2.48)				

Table VIII

Short sale ban sample regression results

Table VIII reports cross-sectional regression results of cumulative return (CUMRET) for stocks included in the SEC short sale ban regressed on an option dummy which equals 1 if the stock is listed on the Chicago Board Option Exchange, the Philadelphia Stock and Option Exchange or the American Stock and Option Exchange and control variables. Cumulative return is calculated from September 22, 2008 to October 8, 2008. INST is the percentage of shares outstanding held by institutional investors on August 30, 2008. SIZE is market capitalization on the day prior to the short sale ban. TURN and ILLIQ are calculated over the year prior to the short sale ban. TURN is the average daily ratio of shares traded divided by shares outstanding. ILLIQ is the average weekly Amihud Illiquidity Ratio (Amihud, 2002) calculated as the absolute weekly return divided by the dollar value of weekly trading volume. BK/MK is book value on 2007 fiscal year end for each stock divided by SIZE. β_{BAN} is the beta coefficient from the regression of daily excess stock return regressed on the excess equal weighted average daily return of the banned stocks over the year prior to the short sale ban.

Dependent Variable	Intercept	OPTION	INST	SIZE	TURN	ILLIQ	BK/MK	β_{BAN}	adj R^2
CUMRET	0.00	**-0.17**	0.048	0.018	-0.003	0.003	0.002	**-0.39**	0.27
	(0.00)	(4.70)	(1.53)	(0.60)	(0.11)	(0.09)	(0.07)	(10.49)	
CUMRET	0.00	**-0.18**						**-0.39**	0.27
	(0.00)	(5.44)						(11.58)	

Figure 1

Institutional ownership at the time of option introduction

The left hand Y axis reports the mean percentage of total shares outstanding held by institutional investors, by quarter, centered on the option listing quarter. The right hand Y axis reports the mean number of institutional investors which have holdings in each stock, in each quarter, over the same time span. List quarter is the quarter of option introduction.

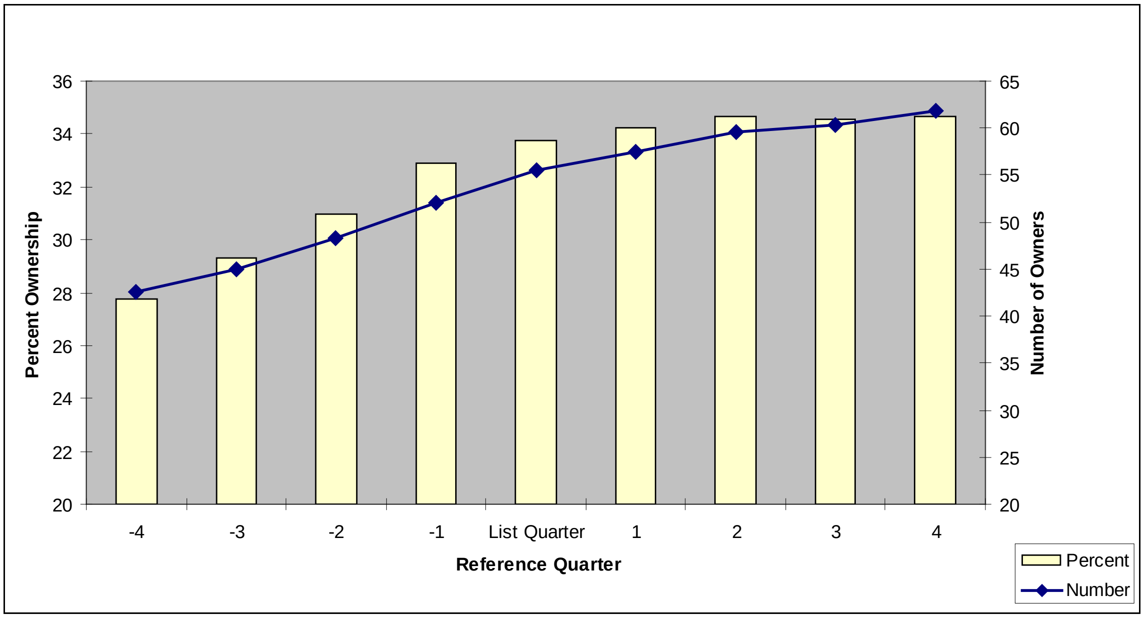
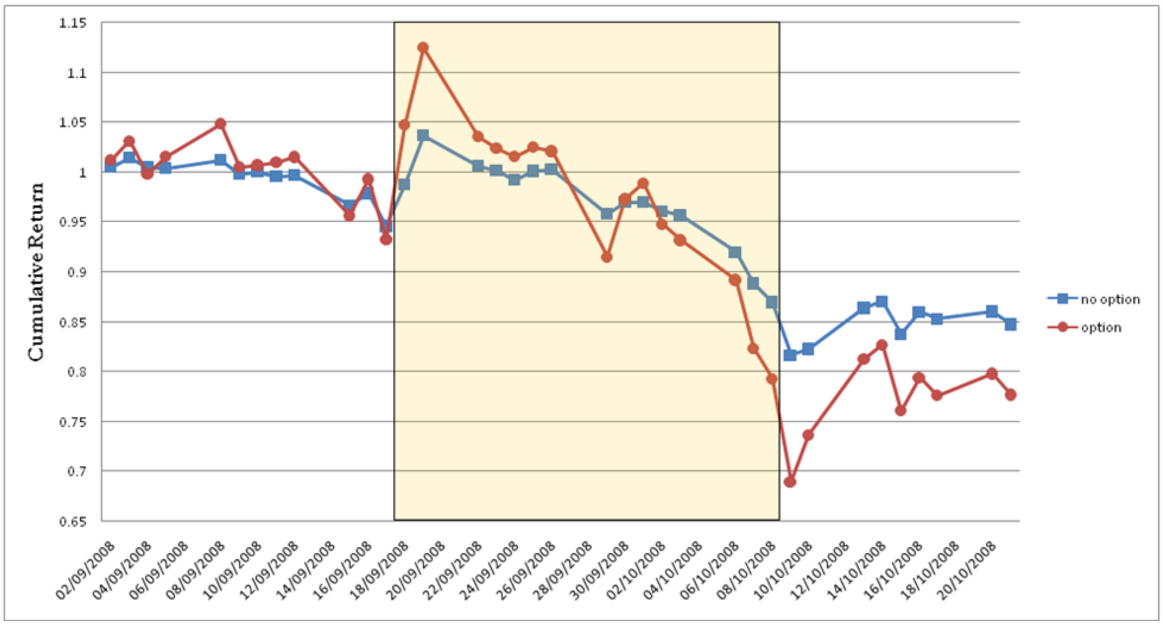


Figure 2

Cumulative returns for optioned and non-optioned stocks during the short sale ban

Figure 2 displays cumulative abnormal returns to finance sector stocks for which short sale transactions were banned from September 19, 2008 to October 8, 2008 (shaded region of the figure). Cumulative returns are reported separately for stocks listed and not listed on the Chicago Board Option Exchange, the Philadelphia Stock and Option Exchange or the American Stock and Option Exchange.



Chapter 3

What Determines the Success of Cross-Listings in the U.S.?

As Karolyi (2006) notes in his extensive survey of the literature on foreign equity listings, we now know a great deal about the valuation and market microstructure effects of cross-listings, especially when these involve stock exchanges in the U.S. What we know less about is the extent of disparity in post cross-listing success, and its determinants. Our paper addresses these issues by studying liquidity and returns for a sample of 179 cross-listings in the U.S. between 1981 and 2004. We first document wide variation in post cross-listing liquidity and returns. We then investigate why some firms see very liquid trading and high prices while others do not. Specifically, we ask whether pre-listing firm and country characteristics predict post-listing performance, and whether U.S. institutions play any part in post-listing success.

Our choice of post-listing liquidity and stock returns as the twin metrics of success is rooted in survey responses by managers of cross-listing firms. As Mittoo (1992) and Fanto and Karmel (1997) report, managers frequently cite access to U.S. investors, improved liquidity and future capital-raising as important goals of cross-listing. A liquid market characterized by low-cost, active trading is an indication of the degree of interest shown by U.S. investors, and a critical variable in successfully tapping the U.S. equity markets for new capital. Managers are keenly interested in higher post-listing prices in the U.S., not least because, consistent with the literature on seasoned equity offerings, capital-raising is easier for a firm enjoying a favorable run.[30]

[30] See, for example, Baker and Wurgler (2000) and Pontiff and Schill (2002).

Our sample of 179 foreign cross-listings in the U.S. between 1981 and 2004 closely corresponds to the universe of cross-listings on the major U.S. exchanges over this period. The sample firms differ significantly in terms of pre-listing characteristics. For instance, we sort the sample into five groups based on the pre-listing cumulative annual return. The lowest and highest quintiles have mean annual returns of -22% and +44%. Moreover, the firms come from 22 countries that differ in terms of governance and financial development. We also see wide variation in U.S. institutional ownership in the cross-listing quarter—forming quintiles on the basis of institutional ownership, the means for the bottom and top quintiles are 0% and 9.5%.

Our analysis focuses on liquidity (captured by the Amihud illiquidity ratio)[31] and cumulative excess returns in the U.S. over the year following cross-listing. Initially, we document that there is considerable dispersion in both liquidity and cumulative returns in the post-listing period. When we sort firms into quintiles based on illiquidity, the ratio of the highest to lowest quintile means is in excess of 10^3. Only the highest quintile firms enjoy liquidity comparable to that of the median NYSE-listed firm. Similarly, forming quintiles on the basis of the post-listing cumulative excess return, we see that the top and bottom quintiles have mean returns of -44% and +128%. Adjacent quintile means and medians for both variables are also significantly different from each other. It is this disparity in success that we seek to explain.

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[31] As discussed in Section II, the Amihud illiquidity ratio (*AIR*) is defined as the average ratio of the absolute weekly return to the dollar value of weekly volume (see Amihud, 2002). Note that the *AIR* is an inverse measure of liquidity.

We find that the post-listing liquidity and valuation benefits of cross-listings are crucially dependent both on prior home-market success and on U.S. institutional holdings in the cross-listing quarter. Stocks with greater institutional ownership upon cross-listing see more liquid U.S. trading. Two plausible explanations are that institutional ownership generates trading, and that institutions are attracted to stocks with minimum acceptable expected post-listing liquidity. Additionally, firms with a higher abnormal price run-up in the year prior to cross-listing and firms that see more liquid domestic trading enjoy greater post-listing liquidity in the U.S. However, post-listing U.S. liquidity is unrelated to the cross-listing firm's home market index performance, underscoring the importance of firm-specific performance in driving post-listing success.

A similar picture emerges when we examine cumulative post-listing returns. Foerster and Karolyi (1999) document the tendency for firms to cross-list after a period of home country price run-ups; further, consistent with market timing, these run-ups are, on average, reversed. A similar finding is documented for cross-listings in other countries in Sarkissian and Schill (2008). We find that whether or not a run-up is sustained depends on aggregate institutional holdings in the cross-listing quarter, and that firms with higher institutional ownership earn higher post-listing excess returns. Thus, firms with high U.S. institutional holdings tend to sustain their pre-listing run-ups, but firms with low U.S. institutional holdings appear to reverse their pre-listing gains.

It is possible that institutional ownership at the time of cross-listing is driven in part by pre-listing firm and home market characteristics, and that the significance of institutional ownership in determining post-listing liquidity and returns is therefore spurious. To address this concern, we model U.S. institutional ownership as a function of pre-listing variables, and use the residual institutional ownership from this model in our analysis. Our conclusions regarding the importance of institutional ownership in cross-listing success are largely unchanged.

This analysis has the added virtue of shedding light on factors that attract U.S. institutional investors to foreign listings. Institutional ownership is positively associated with the pre-listing firm-level price run-up but negatively related to home market index performance. These opposing effects suggest that institutions seek firms with superior idiosyncratic performance and are wary of firms from countries whose markets have been hot. Further, we find that institutions invest in stocks that pay dividends, are large, and trade at relatively low P-E multiples (so-called value stocks). There is a preference for holding direct listings from Canada vis-à-vis ADRs from other countries, consistent with the proximity preference documented in Sarkissian and Shill (2004). Interestingly, institutions appear to invest more heavily in ADRs from countries that are more corrupt. One interpretation of this result is that investors shun stocks from corrupt countries until they are listed on a U.S. exchange. This view lends support to the argument proposed by Stulz (1999) and Reese and Weisbach (2002) that a U.S. cross-listing affords firms a chance to "lease" a superior corporate governance regime, something that is particularly valuable to high quality firms from more corrupt countries. The passage of the Sarbanes-Oxley Act in 2002 does not seem to have had a material impact on institutional holdings of cross-listed stocks, nor on post cross-listing success.

Taken together, our results suggest that cross-listing success depends on two conditions. First, domestic winners tend to succeed in the U.S. Second, firms with high U.S. institutional holdings upon cross-listing generally enjoy greater subsequent success. When these two pieces are in place, firms are more likely to see higher long-term liquidity and prices in the U.S. Our results imply that cross-listing does not offer an unconditional guarantee of success; rather, cross-listing success is best seen as a complement to domestic success.

The remainder of this paper is organized as follows. Section I discusses the related literature in more detail and develops the hypotheses we test. Section II describes the data, Section III presents our results and Section IV concludes.

I. Hypotheses and Related Literature

A large literature deals with the benefits of international cross-listings.[32] Theoretical work (e.g. Errunza and Losq, 1985) has suggested that cross-listing allows firms to circumvent investment barriers and thereby lowers expected returns. In support of this hypothesis, empirical work shows that announcements of cross-listings in the U.S. typically are associated with positive abnormal returns.[33]

[32] Karolyi (1998, 2006) and King and Mittoo (2007) provide excellent surveys.

[33] See, for example, Miller (1999).

Several alternative hypotheses have been advanced to explain why foreign firms realize a net benefit to listing in the U.S. Lins et al. (2005) argue that firms cross-list in the U.S. to get around underdeveloped local capital markets and gain access to the greater liquidity of the U.S. market. They show that cross-listed firms become less credit constrained and that firms with the largest investment opportunities realize the greatest benefits from cross-listing. In the corporate governance literature, Stulz (1999) and others suggest that cross-listings allow firms to "rent" foreign governance and legal systems, for example, by listing in countries where minority shareholder protection rights are stronger. Similarly, Doidge et al. (2001, 2008) hypothesize that cross-listing in the U.S. limits the ability of large controlling shareholders to extract private benefits from the firm. Thus, cross-listed firms gain value as a smaller fraction of their cash flow is expropriated by controlling shareholders. Consistent with this view, firms from weaker corporate governance regimes benefit more from a U.S. listing. While explaining the overall benefits of cross-listing is not the focus of our analysis, we will control for these effects in our tests.

We define success in terms of post-listing liquidity and the permanence of the pre-listing price run-up. Our choice of success metrics is based on survey evidence reported in Mittoo (1992) and Fanto and Karmel (1997). Managers' responses to these surveys indicate that corporate motivations for cross-listing include expansion of the firm's U.S. shareholder base, access to U.S. capital markets, and expansion of its business operations.

Higher liquidity expands the shareholder base, especially that consisting of institutional owners (Ferreira and Matos, 2006), and is associated with a lower cost of capital (Amihud and Mendelson, 1986; Hasbrouck, 2006). Thus, firms would like to see greater liquidity for their cross-listed shares in the U.S. markets. Overall, the liquidity effects of cross-listings appear to be positive. For example, Foerster and Karolyi (1998) document a reduction in the average bid-ask spread for a sample of Canadian stocks listing in the U.S. However, improved liquidity is confined to stocks that experience a significant shift in trading volume to the U.S. exchange. We measure liquidity using the Amihud illiquidity ratio (*AIR*), described in Section II.[34]

Higher post-listing stock prices are desirable because they facilitate equity capital-raising in the U.S. To capture price performance, we examine cumulative excess returns in the U.S. following the cross-listing event. Forester and Karolyi (1999) find that average cumulative abnormal returns are negative over the year following cross-listing, consistent with market timing motivations for cross-listing firms. Our study examines whether and why some firms have higher post-listing cumulative returns than others.

In order to explain cross-sectional variation in *AIR* and post-listing cumulative excess returns (*CUMRET*), these variables are modeled as follows:

[34] Several recent papers have examined the volume effects of cross-listings, e.g. Domowitz, Glen and Madhavan (1998) and Karolyi (2003). Baruch, Karolyi and Lemmon (2008) study cross-sectional variation in the U.S. share of trading and Halling, Pagano, Randl and Zechner (2008) examine trading volume in the U.S. and in the domestic market following cross-listing. We examine cross-sectional variation in liquidity (which is related to, but broader than, trading volume). We also examine cross-sectional variation in share turnover. Since we reach the same conclusions for turnover and liquidity, we only tabulate and discuss the results for the latter.

$$AIR_i = a + \sum_{j=1}^{J} b_j . B_i^j + \sum_{j=1}^{J} c_j . C_i^j + d_1 INST_i + d_2 \Delta INST_i + e_i \qquad (3.1a)$$

$$CUMRET_i = a + \sum_{j=1}^{J} b_j . B_i^j + \sum_{j=1}^{J} c_j . C_i^j + d_1 INST_i + d_2 \Delta INST_i + e_i \qquad (3.1b)$$

AIR is measured weekly between the second and fourth quarters (i.e. between 70 and 250 trading days) after the cross-listing date, while *CUMRET* is based on daily data from the U.S. and measured over the same interval.[35] *INST* is the fraction of shares held by U.S. institutional investors (drawn from 13F filings) at the end of the quarter containing the cross-listing date for firm *i*. *ΔINST* is the change in institutional ownership from the end of the cross-listing quarter through the end of the fourth quarter after cross-listing. The explanatory variables (described below) are divided into two categories: B_i^j are firm-specific variables, while C_i^j are country-level variables.

As shown in (3.1a) and (3.1b), we assess the effects of U.S. institutional ownership (*INST*) on illiquidity and cumulative returns. Note that the effects of *INST* will be incremental to those of the other independent variables. We conjecture that firms that are able to attract greater U.S. institutional ownership will realize greater cross-listing success (i.e. lower values of *AIR* and higher values of *CUMRET*).

[35] Our results are robust to measurement of *AIR* and *CUMRET* over the window of [0,250] trading days following cross-listing. We focus on the window of [70,250] trading days following cross-listing for several reasons. First, excluding the quarter following cross-listing minimizes the potential for abnormal trading patterns associated with cross-listing to bias the measurement of *AIR*. Second, measuring both *AIR* and *CUMRET* one quarter following cross-listing ensures that institutional ownership is pre-determined relative to the measurement window. As Thomson Financial tracks institutional ownership on a quarterly basis, the first available ownership observation may lag the cross-listing date by up to three months, depending on the proximity of the cross-listing date to the end of the quarter.

This conjecture is reasonable on the premise that institutional investors are deemed sophisticated, generally possess higher quality information and, by virtue of their experience, are able to pick better-performing stocks, at least vis-à-vis smaller investors in the U.S. It is also consistent with the inferences in Foerster and Karolyi (1999) regarding the role of institutions in the cross-listing announcement and listing date stock returns, and with investor recognition effects identified in Merton (1987). Note that we have included the change in institutional ownership (*ΔINST*) as an additional control variable in the model. To the extent that liquidity or price changes are associated with contemporaneous institutional ownership changes (e.g. because increased institutional holdings contribute to higher liquidity or prices), a model that excludes this variable suffers from an omitted variables bias.

A concern with estimating (3.1a) and (3.1b) is that institutional ownership and, correspondingly, the change in institutional ownership might not be exogenous in relation to expectations of post-listing liquidity or post-listing returns. We address the potential endogeneity of institutional ownership by first modeling *INST* as a function of pre-listing firm and market characteristics in (3.1c) below, and then using the residuals from this model in (3.1a) and (3.1b). To the variables considered as potential determinants of *AIR* and *CUMRET* (discussed below), we add two variables as potential determinants of *INST*:

1. Analyst following (*ANAL*). The number of analysts following the cross-listed firm is measured at the end of the quarter that includes the cross-listing date. If analysts increase information availability and

transparency (e.g. Brennan and Subrahmanyam, 1995), institutions are likely to increase their holdings, and the coefficient for institutional ownership should be positive.

2. The firm's price-earnings ratio (*PE*), measured as the ratio of the stock price as of a month before the cross-listing date divided by annual earnings at the end of the fiscal year before the cross-listing event. To the extent that firms with higher P-E ratios are believed to be overvalued, institutions should avoid them. The coefficient for ownership is thus expected to be negative.

Adding in the firm-specific (B_i^j) and country-specific (C_i^j) determinants of institutional ownership, *INST* is modeled as follows:

$$INST_i = a + \sum_{j=1}^{J} b_j . B_i^j + \sum_{j=1}^{J} c_j . C_i^j + hANAL_i + jPE_i + e_i \tag{3.1c}$$

Next, we describe the B_i^j and C_i^j and predictions regarding their coefficients. Some of these variables potentially affect liquidity or returns directly; others might do so only indirectly through their effects on *INST*. We include the following six firm-specific variables:

1. The firm's size in the home market (*SIZE*). This is measured as the firm's market capitalization in U.S. dollars as of one month before the cross-listing date. If larger firms are viewed as more stable, institutions will buy more shares in larger firms. Thus, for institutional ownership in (3.1c), the coefficient on size is expected to be positive. If information asymmetry is less

marked for larger firms, they are more likely to have greater liquidity in the U.S. market.[36] We do not have a prediction for the relation between post-listing cumulative returns and size.[37]

2. The firm's relative size (*RELSIZE*). This variable is measured as the natural logarithm of the ratio of the firm's market capitalization to aggregate market capitalization one month before the cross-listing date.[38] If institutions follow value-weighted indexes they will buy more shares in firms that constitute a larger fraction of home country capitalization. Moreover, larger firms are likely to be among the more successful and visible offerings from the country. Consequently, the coefficient on *RELSIZE* is expected to be positive for institutional ownership. As with *SIZE*, we expect *RELSIZE* to be positively associated with post-listing liquidity. We have no prediction for the relation between post-listing cumulative returns and *RELSIZE*.

3. The firm's book-to-market ratio *(BM*). *BM*, measured using data from the fiscal year preceding the cross-listing date, is a measure of growth opportunities, with lower values of BM belonging to growth firms.

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[36] Substantial supportive evidence exists in the corporate finance and accounting literature. For instance, Zeghal (1984) shows that the reaction to financial releases is lower for large firms, presumably because much of the information has already been impounded in prices.

[37] It is possible that larger firms are less subject to market timing (e.g. due to lower levels of information asymmetry), and therefore less likely to display price reversals. In this event, the predicted coefficient on size will be positive.

[38] A log transformation is used because a handful of firms account for a very large proportion of total market capitalization in some emerging markets (see Table VIV).

The sign of the coefficient in the institutional ownership regression depends on whether institutional investors prefer, at least in our sample period, growth or value exposure. We do not have clear predictions for the coefficients in the illiquidity and return regressions.

4. Pre-listing illiquidity (*ILLIQ*). Assuming that institutions have a preference for holding liquid assets, the coefficient for ownership is expected to be negative. If there is persistence in liquidity, home market illiquidity will (positively) predict U.S. illiquidity. As mentioned above, greater illiquidity should be reflected in higher expected returns (e.g. Amihud and Mendelson, 1986). Thus, U.S. cumulative returns should be positively related to home market illiquidity.

5. Dividend dummy (*DIV*). This is a dummy variable corresponding to whether the firm has paid a dividend in the year prior to cross-listing. Firms that pay dividends are expected to be more stable and thus more attractive to institutions; thus, the coefficient for ownership is expected to be positive. We do not have a definite prediction on whether dividends are associated with higher post-listing returns. To the extent that dividends are associated with greater stability, we expect a negative coefficient for illiquidity.

6. The firm's cumulative excess return (*RUNUP*) over the year preceding the cross-listing date.[39] There are at least two possible interpretations of this variable. If larger run-ups are symptomatic of more severe overvaluation, institutions should avoid stocks with larger run-ups. On the other hand, if run-ups imply superior performance and are a signal of quality, institutional ownership should be higher. Thus, the relation between *INST* and *RUNUP* is an empirical issue. The coefficient for U.S. cumulative returns is expected to be negative if run-ups imply overvaluation. We do not postulate a definite sign for the relation with post-listing liquidity (though there is the possibility that investors may be attracted to firms that have had a great domestic run).

To this set of firm-level variables, we add the following four country-level variables:

1. The country's corruption index (*CORRUPT*) for 2007 is obtained from Transparency International.[40] If firms from more corrupt countries (lower values of *CORRUPT*) are likely to be shunned by U.S. institutions, *INST* will be lower. On the other hand, institutional ownership could be high if U.S. listings provide a governance benefit. In particular, it is possible that the only firms able to cross-list in the U.S. from corrupt

[39] While computing all pre-listing firm-specific and market returns, we omit the month before the cross-listing date. Thus, the cumulative return measurement window ends one month before the cross-listing date.

[40] Transparency International (http://www.transparency.org/) uses both expert opinions as well as public surveys to assess the level of corruption in a country.

countries are high quality firms that are able to commit to better governance. By virtue of this self-selection, such firms will be attractive to institutions. Thus, the coefficient for *INST* is ambiguous. The same arguments suggest that the coefficient for illiquidity is ambiguous. We have no definite prediction for post-listing cumulative returns.

2. The cumulative home market index return over the year preceding the cross-listing date (*HCINDX)*. If firms cross-list on the heels of high home market returns (consistent with market timing objectives), U.S. institutions should avoid such offerings and the coefficient for ownership will be negative. The coefficients for illiquidity and cumulative returns will depend on whether U.S. investors condition on prior foreign country index performance and whether they chase foreign performance or view it with suspicion. Hence, we cannot offer definite predictions for the liquidity or cumulative return coefficients.

3. The cumulative U.S. market index return over the year preceding the cross-listing date (*USINDX*). U.S. institutions are expected to be net buyers if U.S. market conditions are buoyant. If firms cross-list in these circumstances, institutional ownership should be higher and the coefficient for *INST* will be positive. Liquidity ought to benefit as well, although we do not have a prediction for post-listing cumulative returns.

4. A dummy for Canadian listings (*CAN*). There are three effects that separate listings from Canada and other countries: Canadian firms offer direct listings rather than ADRs; Canadian stocks are likely to be more

familiar to U.S. investors, and Canadian and U.S. stocks trade in the same time zone. If these effects serve to increase U.S. investor interest, the coefficient on *CAN* will be positive. For the same reasons, Canadian stocks are predicted to enjoy greater post-listing liquidity. No prediction is offered for post-listing cumulative returns.

The determinants of institutional ownership (3.1c) are discussed in *sub-section a* and we discuss the estimates from (3.1a) and (3.1b) in *sub-sections b and c*, respectively. All equations are estimated with a Sarbanes-Oxley dummy variable (SOX) set equal to 1 for cross-listings occurring during and after 2002, and zero otherwise.

II. Data

We identify the set of all non-U.S. firms that list their shares on the New York Stock Exchange (NYSE), American Stock Exchange (AMEX) or NASDAQ between 1981 and 2004. Foerster and Karolyi (1999) find abnormal price declines for foreign stocks listing in the U.S. over a one-year window following the listing. Accordingly, we use a one-year post-listing window to evaluate stock success. We also use a one-year window prior to cross-listing to measure pre-listing performance.

Price, volume and shares outstanding data are collected from DataStream. If post cross-listing data are missing from DataStream the data are obtained from the

Center for Research in Security Prices (CRSP) tapes. Stocks missing one year of data prior to cross-listing or one year of data after cross-listing are excluded. Thus, stocks listing in the U.S. within one year of listing on the domestic exchange or stocks initially listing in the U.S. are excluded. Book value data are obtained for the fiscal year prior to cross-listing. Book value data for this year could not be located for 17 stocks. For these stocks, book value is collected for the closest preceding year.

Institutional ownership data are obtained from the Shareworld 13F filing database maintained by Thomson Financial. The Shareworld 13F database tracks, by quarter, the share holdings of institutional investors based on 13F filings made with the U.S. Securities and Exchange Commission. Any manager with more than $100 million at their discretion is required to make a quarterly filing of a 13F for every security holding in excess of $200,000 or 10,000 shares. The institutions represented in Shareworld include mutual funds, banks, insurance firms and pension funds. At the time of dataset construction, the Shareworld database tracked 13F filings from 1980 through 2005. To ensure the availability of one year of post-listing ownership data, all cross-listings after December 2004 are excluded. Institutional ownership data are recorded at the end of the quarter when cross-listing occurs and four quarters thereafter.

A total of 179 stocks meet the data requirements. Panel A of Table VIV summarizes the number of cross-listings by country of origin. Twenty-two countries are represented. The majority of listings originate from Canada (83), followed by developed European countries (53), Asian countries (15) and Australia (10). On average, cross-listings from emerging markets have a higher market capitalization relative to total home market capitalization than do firms from developed countries. Firms from countries rated as more corrupt (with lower values of Transparency International's corruption index) also appear to be larger when measured relative to aggregate home market capitalization. One interpretation is that smaller firms from emerging markets and more corrupt countries find it more costly to cross-list in the U.S. because they are not well-known.

In Panel B of Table VIV we provide statistics for the cross-listed firms prior to, and at the time of, the U.S. listing. On the date of cross-listing the mean and median home country stock prices, in U.S. dollars, are \$19.29 and \$16.00, and the firms have a mean and median market capitalization of \$6.5 billion and \$1.1 billion. Daily trading volume on the domestic exchange over the year prior to cross-listing averages about 1.3 million shares, and the average daily turnover is 0.63% of shares outstanding.

The Amihud Illiquidity Ratio (*AIR*) is analogous to the price impact of dollar volume. It is calculated as the mean value of the absolute weekly return in U.S. dollars divided by the U.S. dollar value of weekly trading volume, i.e. $AIR = \frac{|r_w|}{Volume_w}$, where r_w is the close-close return in week w and $Volume_w$ is the week w volume.[41]

[41] The *AIR* is calculated using weekly data to overcome problems related to days with no trading volume.

For each stock, we calculate *AIR* for the year preceding the cross-listing date and report the cross-sectional average. The median statistic is 1.22E-08 and implies that for more than half of the stocks the price impact of trades is small.[42] The median cross-listed firm pays a modest annual dividend ($0.061 per share), and reports an EPS of $0.82.

III. Results

As described above, our analysis focuses on two post-listing success measures. The first is liquidity (based on the *AIR*). In addition, we study the post-listing cumulative excess return (*CUMRET*) as a measure of the cross-listing firm's price performance in the U.S. *CUMRET* is calculated as the compounded value of the daily market-adjusted return between day 70 and day 250 relative to the cross-listing date (which is set to be day 0),

$$CUMRET_i = \prod_{t=70}^{250} (1 + R_{i,t} - R_{m,t}) \tag{3.2}$$

where $R_{i,t}$ is the return for stock i on date t and $R_{m,t}$ is the CRSP equal-weighted market return, also on date t.

We start by assessing the extent of cross-sectional variation in *AIR* and *CUMRET* after the cross-listing event. In Figure 3A, we divide the sample firms into five equally populated bins ranked by liquidity, and plot the logarithm (to the base

[42] The mean statistic is influenced by a few very large observations.

10) of the mean *AIR* for each bin (the log transformation allows us to fit the values into the same figure). The log mean *AIR* varies from -8.4 to -5.2, a difference of three orders of magnitude from the top to the bottom quintile, and the difference from one quintile to the next is pronounced. Thus, there is wide variation in liquidity for firms in the twelve months following a U.S. listing.

In Figure 3B we similarly partition the sample firms into *CUMRET* quintiles and plot the mean post-listing *CUMRET* for each quintile (calculated over days [70,250] relative to the listing date). The mean *CUMRET* varies from a low of -44% to a high of +128%, and, as with *AIR*, the mean *CUMRET* for each quintile is noticeably different from the means for the adjacent quintiles. While these variations exist by construction, they demonstrate that the cross-listed firms have substantially different experiences in the post-listing period.[43] In the following sections, we implement cross-sectional regressions to uncover the determinants of post-listing liquidity and excess returns. First, we study institutional ownership.

a. Determinants of U.S. Institutional Ownership

There are two reasons for being interested in the determinants of U.S. institutional ownership. The first, described in the previous section, is associated with econometric concerns. The second motivation stems from economic

[43] Given the relatively long sample period, we must address the concern that we are ignoring secular effects. We regress the *AIR* or *CUMRET* for each firm on quintile dummies and year dummies that provide a control for year-specific effects. The coefficients on the quintile dummies are significantly different from each other. Thus, controlling for year effects, illiquidity and price performance continue to vary significantly across the quintiles.

considerations. Specifically, there are several reasons why firms that cross-list their shares in the U.S. would like to attract institutional shareholders. For example, having an institutional shareholder base may be desirable from the point of view of issuing shares in the U.S. in the future. Institutional demand may also help support prices under the assumption that demand curves for the cross-listed shares are downward-sloping. Naturally, then, firms should be interested in knowing the determinants of institutional holdings of cross-listed shares.

Using a rich dataset, Ammer et al. (2005) examine the determinants of overall U.S. ownership, focusing on accounting measures and country-level investor protection. They find that improved accounting practices associated with cross-listing contribute to increases in U.S. ownership; that stocks from English speaking countries and stocks included in the MSCI World Index realize smaller increases in ownership; and that cross-listing has little effect on U.S. holdings of stocks from countries with weak investor protection. We augment their analysis by focusing on firm-level returns and liquidity while also investigating the significance of country-level transparency and corruption. A key difference between the Ammer et al. (2005) study and ours is in the definition of U.S. ownership: Ammer et al. (2005) measure total U.S. ownership; we focus on shares held by U.S. institutions. Our focus on institutional holdings addresses a somewhat different set of questions.

As discussed in Section 1, we model *INST* as a function of past firm-specific and country-specific variables. Coefficient estimates are presented in Table X, with Model 1 containing the full specification and Model 2 including only variables found to be significant at the 10% level of significance. Table X provides evidence of a strong relation between post-listing U.S. institutional ownership and prior stock performance. U.S. institutional ownership tends to be higher for stocks that realize high cumulative excess returns over the year preceding the cross-listing (*RUNUP,* p-value<0.01). Interestingly, institutional ownership is negatively associated with domestic market index performance (*HCINDX*), suggesting that institutions are not chasing foreign country-level performance; rather, they buy more shares in stocks that are superior performers relative to their markets. We also find that stocks that pay dividends (*DIV*), have a larger market capitalization (*SIZE*) and a smaller P-E ratio (*PE*) realize higher U.S. institutional ownership, indicating a preference for more conservative stocks among institutional investors.

Institutional holdings are not significantly related to relative market capitalization (*RELSIZE*), book-to-market (*BM*), pre-listing domestic liquidity (*ILLIQ*) and pre-listing U.S. index returns (*USINDX*), nor to the post-listing presence of analysts (*ANAL*).[44] We find a negative relation between U.S. institutional ownership and the corruption index for a country (*CORRUPT*), so U.S. institutional ownership at the time of cross-listing is higher for firms from more corrupt countries.[45] (Recall that the corruption index is inversely related to the level of corruption in a country.) This could happen if U.S. institutions stay away from corrupt country stocks *until*

[44] Dahlquist and Robertsson (2001) study foreign ownership of Swedish firms and find that foreigners prefer larger firms, low dividend-yield firms, and tend to be institutions rather than individual investors.

[45] As an alternative to the corruption index, we also use an indicator variable for emerging vs. developed economies. Since the results are materially similar, we only report and tabulate the results with the corruption index.

they list in the U.S., and is consistent with the evidence in Ammer et al. (2005) that cross-listings are associated with a large increase in U.S. shareholdings. Finally, as seen in the positive coefficient on *CAN*, U.S. institutions favor Canadian listings relative to ADRs. A plausible explanation is that U.S. investors are more familiar with Canadian firms due to the geographic proximity of the two countries and thus hold larger proportions of Canadian stocks.[46] The SOX dummy is not significant, indicating little effect of tighter post-Sarbanes-Oxley regulations on institutional holdings.

b. Liquidity

As mentioned in Section I, liquidity on the U.S. exchange is a direct measure of cross-listing success. Post-listing liquidity is measured inversely as the Amihud Illiquidity Ratio (*AIR*), described in Section II. We analyze the cross-sectional determinants of the *AIR* by estimating (3a) for the mean weekly *AIR* in U.S. trading. The weekly *AIR* in the U.S. is computed between day 70 (the end of the first quarter) and day 250 (the end of the fourth quarter) following the cross-listing date.[47] All specifications employ industry fixed effects, defined at the level of the 2 digit SIC code.[48] The coefficients are reported in Table XI. Model 1 and model 3 report the full model specification using *INST* and *Residual INST*, respectively. Model 2 and

[46] Sarkissian and Schill (2004) document a pronounced cultural and geographical proximity bias in their study of foreign market choice for cross-listing firms.

[47] The cross-sectional median of the firm-level *AIR* in U.S. trading is 6.5E-8, similar to the median of 7.1E-8 reported in Hasbrouck (2006) for a sample of U.S. firms.

[48] Including country and year fixed effects does not materially alter our results.

model 4 report the reduced model specifications, including only variables significant at the 10% level of significance.

AIR has a negative coefficient on *Residual INST* (p-value<0.01). Thus, liquidity is positively related to residual institutional ownership, suggestive of a beneficial effect of U.S. institutions on liquidity. One explanation for this relation is that increased institutional ownership boosts trading and thereby promotes liquidity. Recall that this measure of institutional ownership is cleansed of pre-listing influences on the level of U.S. institutional ownership.[49]

Firms with larger scaled market capitalization (*RELSIZE*) see lower values of *AIR*, i.e. are more liquid in U.S. trading. By contrast, the coefficient on market capitalization (*SIZE*) is not significant. Thus, the absolute size of the firm is not important in driving liquidity; rather, what matters is how large the firm is relative to other firms from its home market. It seems that firms that are relatively large at home are more likely to attract trading interest in the U.S. Thus, for two firms with the same dollar market capitalization, the firm from the smaller market is more likely to experience greater liquidity in the post cross-listing period.

The negative coefficient on *RUNUP* means that firms enjoying larger cumulative excess returns in the 12 months prior to cross-listing experience higher liquidity after cross-listing. This result suggests that prior home market success engenders greater post-listing liquidity. Book to market (*BM*), the corruption index

[49] When we use the actual level of institutional ownership instead of the residuals from table X (compare model 1 and model 3, for instance), the coefficient and its significance level are virtually identical, suggesting that concerns regarding the endogeneity of institutional ownership are probably not important.

(*CORRUPT*), and the U.S. and domestic market returns (*USINDX* and *HCINDX*) are not significantly associated with post-listing liquidity. The coefficient on the SOX dummy is also not statistically significant. Somewhat surprisingly, the coefficient on the Canada dummy (*CAN*) is positive and significant in two models. Thus, all else constant, Canadian stocks are less liquid in U.S. trading than other foreign issues.

Overall, the results point to a positive effect of U.S. institutions on post-cross-listing U.S. liquidity. Controlling for U.S. institutional interest, firms that are large in their home markets and firms that have performed well in terms of delivering high excess returns are most likely to enjoy deeper markets in the U.S. These results suggest that performance thresholds promote investor interest and liquidity. In particular, if institutions herd on the basis of the same signal such as pre-listing performance, their buying decisions will be correlated. Greater institutional ownership will eventually show up as greater two-sided liquidity, as some institutions will always be selling.

c. Post Listing Excess Returns

In a manner consistent with Foerster and Karolyi (1999), we start by examining the cumulative market-adjusted return (*CUMRET*) on the domestic exchange the year before and after cross-listing. Figure 4 displays the mean *CUMRET* for the full sample and for terciles sorted by U.S. institutional ownership as of the cross-listing quarter. On average, across the full sample, stocks experience a

15% cumulative abnormal run-up in price over the year preceding the cross-listing date and a cumulative abnormal return of -10% over the year after the listing.[50]

More interesting is the variation in *CUMRET* as a function of U.S. institutional ownership. The pre-cross-listing run up is 36% for the high institutional ownership tercile compared to 0.09% for the low ownership tercile. Additionally, the post-listing *CUMRET* for the high institutional ownership tercile is +6% compared to -6% and -32% for the low and moderate ownership terciles.[51] Thus, on average, high institutional ownership stocks see higher pre-listing price run-ups but are able to avoid the post-listing declines experienced by moderate and low ownership stocks. This suggests that greater institutional ownership in the cross-listing quarter is associated with higher stock prices in the subsequent three quarters.

Table XII presents the results of the cross-sectional regression analysis of the post-listing cumulative excess return, described in (3.1b). *CUMRET* is measured from day 70 through day 250 after cross-listing. The independent variables are described in Section 1. Replicating the model sequence in the analysis of *AIR*, model 1 and model 3 report the full specifications using *INST* and *Residual INST*, respectively. Model 2 and model 4 report the reduced specifications, including only variables significant at the 10% level of significance.[52]

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[50] Our full-sample *CUMRET* patterns are similar to Foerster and Karolyi (1999), who document a pre-listing run-up of 19% and an average post listing decline of 14% for 153 ordinary and ADR listings. For this analysis, we use returns over the entire period, not just the [70,250] day window. It has the added value of showing returns immediately after the cross-listing date.

[51] This non-monotonicity in ownership disappears in our regressions, once we control for other factors.

[52] Given our specific interest in the effects of institutional ownership, we include this variable in every specification regardless of the associated p-value.

We find that the post-listing *CUMRET* is negatively related (p-value < 0.01) to the pre-listing home market index return (*HCINDX*), consistent with market timing on the basis of general home market conditions. That is, firms that list in the U.S. when the home market is hot display relatively weak performance subsequent to the listing. By contrast, the coefficient on *RUNUP*, the firm-specific pre-listing cumulative return, is positive. This is consistent with the evidence in Sarkissian and Schill (2008) which documents that high Q ratio firms tend to maintain their high Q ratios after they cross-list in a new country. Putting the coefficients on *HCINDX* and *RUNUP* together, it appears that firms from hot markets underperform in the U.S., while firms that display superior abnormal performance relative to their home market continue to do well. This is consistent with *RUNUP* reflecting firm quality rather than market timing. The coefficients also have a message for exchanges. Admitting firms from hot markets actually impedes future success; rather, success is positively related to firm-specific performance in the home market.

The post cross-listing *CUMRET* is positively related to U.S. institutional ownership at the time of the cross-listing. The coefficient on institutional ownership in model 1 and model 2 is on the same order of magnitude as the coefficient on residual institutional ownership in model 3 and model 4 and is usually significant at better than the 10% level of significance. The positive effect of *INST* is consistent with institutional investors propping up share prices for cross-listing firms, and with the evidence in King and Segal (2008) for a sample of Canadian listings in the U.S. Possible explanations include downward-sloping demand curves for stocks or the fact that institutions, as sophisticated investors, are able to identify the better offerings and shun the underperformers.

As with the liquidity regressions in Table XI, however, we find no relation between the post cross-listing *CUMRET* and market capitalization (*SIZE*), illiquidity (*ILLIQ*), the U.S. market return (*USINDX*), the corruption index (*CORRUPT*) or the Canada dummy (*CAN*). There appears to be no significant effect of SOX on post-listing returns. *RELSIZE* has a coefficient that is negative and at least marginally significant, implying that firms that are large in their home markets are associated with lower post cross-listing cumulative returns. This result stands in contrast to the strong positive effect of *RELSIZE* on liquidity, documented earlier.

Overall, the *CUMRET* results support the hypothesis that prior home market firm performance is important in predicting post-listing success. Additionally, institutional ownership appears to positively predict future cumulative returns.

IV. Concluding Comments

In this paper we examine the determinants of success for cross-listings on a U.S. exchange. We measure post-listing success in terms of liquidity, measured by the Amihud illiquidity ratio, and price performance, captured by the cumulative return, and focus on the part played by prior performance and by institutional ownership. As part of our analysis, we also shed light on firm and country characteristics that attract or repel U.S. institutional investors.

Using a sample of 179 foreign listings between 1981 and 2004, we find that pre-listing domestic performance is a good indicator of continued success in the U.S., in terms of both achieving greater liquidity and sustaining the pre-listing home market price run-up. High firm-specific home country returns predict high liquidity and high cumulative returns in the U.S. By contrast, the run-up in the domestic market index is negatively related to post-listing cumulative returns. Firms that are large in their home country tend to see higher U.S. liquidity, though not higher returns. The presence of U.S. institutions at the time of cross-listing appears to have an additional positive influence on liquidity and returns. Studying the determinants of U.S. institutional ownership, we find that institutions are attracted to large, stable, value-type firms that have performed well in their domestic market prior to cross-listing.

These findings point to a market for cross-listings where prior idiosyncratic domestic success is important in achieving the desired goals of market liquidity, high stock prices, and institutional backing in the U.S. Our results imply that cross-listing does not offer an unconditional guarantee of success; rather, cross-listing success is best seen as a complement to domestic success.

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Table VIV

Descriptive and Summary Statistics for Cross-listed Stocks

The cross-listing sample consists of 179 foreign listings in the U.S. originating from 22 different countries between May 1981 and September 2004. Corruption index is the 2007 Corruption Perception Index as calculated by Transparency International. Scaled market cap is the market capitalization for the cross listing stock one month prior to cross listing divided by the total value of its domestic exchange. Price, shares outstanding, and market capitalization are measured one month prior to cross-listing. Volume is the mean number of shares traded per day over the year before cross-listing. Turnover is volume scaled by the total number of outstanding shares. The Amihud Illiquidity Ratio (Amihud, 2002) is defined as the mean of the absolute value of the weekly return scaled by the mean dollar value of weekly trading over the year before cross-listing. Earnings per share (excluding extraordinary items), dividends per share and book value per share are measured at the end of the fiscal year prior to cross-listing. Where applicable, values are stated in U.S. dollars.

Panel A: Summary of the Number of U.S. Cross-listings by Country

Country	N	Corruption Index	Median Scaled Market Cap.	Country	N	Corruption Index	Median Scaled Market Cap.
Argentina	1	2.9	0.00142	Italy	1	4.9	0.003710
Australia	10	8.7	0.0125	Japan	9	7.6	0.001764
Belgium	1	7.3	0.0185	Mexico	2	3.3	0.03447
Brazil	2	3.3	0.0347	Netherlands	6	8.7	0.00386
Canada	83	8.5	0.000700	Russia	2	2.5	0.0879
Finland	1	9.6	0.00080	South Africa	6	4.6	0.0144
France	9	7.4	0.00262	South Korea	4	5.1	0.0254
Germany	8	8.0	0.02447	Sweden	1	9.2	0.00128
Hong Kong	1	8.3	0.0070	Switzerland	4	9.1	0.00245
India	4	3.3	0.0119	Taiwan	1	5.9	0.00507
Israel	1	5.9	0.0151	United Kingdom	22	8.6	0.00365

Panel B: Descriptive Statistics

Variable	Mean	1st Quartile	Median	3rd Quartile
Price ($)	19.29	9.20	16.00	26.00
Shares Outstanding (MM)	227	22	78	220
Market Capitalization ($MM)	6,519	192	1,142	4,362
Volume (1000)	1,301	73	344	1,332
Turnover (%)	0.63	0.16	0.31	0.57
Amihud Illiquidity Ratio ($\times 10^{-6}$)	0.621	0.00183	0.0122	0.133
Earnings per Share ($)	1.36	0.060	0.82	1.75
Dividends per share ($)	0.32	0.00	0.061	0.39

Table X

Determinants of U.S. Institutional Ownership

Table X reports the results of regressions that examine the determinants of U.S. ownership in the quarter foreign firms list in the U.S. The dependent variable is the natural logarithm of shares held by U.S. institutions expressed as a percentage of all shares (*INST* in the text). The independent variables are as follows: *SIZE* is market capitalization one month prior to cross-listing; *RELSIZE* is *SIZE* divided by total market capitalization on the home exchange; *BM* is measured as book value in the year prior to cross-listing divided by *SIZE*; *ILLIQ* is the mean of the absolute value of the weekly return scaled by the dollar value of weekly trading over the year prior to cross-listing; *DIV* is equal to 1 if the firm paid dividends in the year before cross-listing and 0 otherwise; *CORRUPT* is the 2007 Corruption Perception Index as calculated by Transparency International; *ANAL* is the number of analysts following the stock in the quarter of cross-listing; *PE* is the ratio of the price one month prior to cross-listing divided by earnings in the year before cross-listing; *RUNUP* is the compounded daily excess return on the domestic exchange over the year prior to cross-listing; *HCINCX* and *USINDX* are the compounded daily returns in the home country and the U.S. over the year preceding cross-listing; *CAN* is a dummy variable equal to 1 if the cross-listing originates from Canada and 0 otherwise; *SOX* is a dummy variable equal to 1 if the stock cross-lists in 2002 or later. Industry fixed effects are defined on the basis of two digit SIC code. T-statistics are reported in parentheses and coefficients significant at the 10% level of significance appear in bold face. Model 1 presents the full specification and Model 2 retains variables from Model 1 significant at the 10% level of significance.

	Model 1	*Model 2*
Intercept	**-7.11** (3.53)	**-6.71** (6.10)
SIZE	**2.35E-08** (2.10)	**2.23E-08** (2.38)
RELSIZE	-0.088 (0.09)	
BM	-0.0054 (0.61)	
Log ILLIQ	-0.0020 (0.26)	
DIV	**1.01** (2.60)	**1.15** (3.19)
CORRUPT	**-0.28** (2.05)	**-0.268** (2.22)
ANAL	0.046 (1.14)	
PE	**-0.0025** (2.17)	**-0.0026** (2.35)
RUNUP	**0.89** (2.51)	**0.82** (2.41)
HCINDX	**-1.84** (1.69)	**-1.57** (1.93)
USINDX	0.56 (0.47)	
CAN	**4.95** (8.71)	**5.28** (12.80)
SOX	0.031 (0.06)	
Industry Fixed Effects	Yes	Yes
Adj. R^2	0.53	0.55

Table XI

Determinants of Foreign Cross-listing Liquidity

Table XI reports regression results that examine the determinants of liquidity for foreign firms listing in the U.S. The dependent variable is the natural logarithm of illiquidity (*ILLIQ*), defined as the mean of the absolute value of the weekly return scaled by the dollar value of weekly trading. *ILLIQ* is calculated between day 70 and day 250 following cross-listing. The independent variables are as follows: *INST* is the percentage of shares outstanding held by U.S. institutional investors; *ΔINST* is the change in U.S. ownership between the listing quarter and four quarters after the listing date; *Residual INST* is the residual institutional ownership, calculated from (3.1c); *SIZE* is market capitalization one month prior to cross-listing; *RELSIZE* is *SIZE* divided by total market capitalization on the home exchange; *BM* is measured as book value in the year prior to cross-listing divided by *SIZE*; *DIV* is equal to 1 if the firm pays dividends in the year before cross-listing and 0 otherwise; *CORRUPT* is the 2007 Corruption Perception Index as calculated by Transparency International; *PE* is the ratio of the price one month prior to cross-listing divided by earnings in the year before cross-listing; *RUNUP* is the compounded daily excess return on the domestic exchange over the year prior to cross-listing; *HCINCX* and *USINDX* are the compounded daily returns in the home country and the U.S. over the year preceding cross-listing; *CAN* is a dummy variable equal to 1 if the cross-listing originates from Canada and 0 otherwise; *SOX* is a dummy variable equal to 1 if the stock cross-lists in 2002 or later. Industry fixed effects are defined on the basis of two digit SIC code. T-statistics are reported in parentheses and coefficients significant at the 10% level of significance appear in bold face. Model 1 (Model 3) presents the full specification using *INST* (*Residual INST*) and Model 2 (Model 4) retain variables from Model 1 (Model 3) significant at the 10% level of significance.

Independent Variable	**Log U.S. Illiquidity**			
	Model 1	***Model 2***	***Model 3***	***Model 4***
Intercept	**-18.44** (10.55)	**-21.46** (34.17)	**-15.78** (9.16)	**-18.53** (28.74)
Log INST	**-0.34** (5.26)	**-0.35** (5.57)		
Residual INST			**-0.33** (4.59)	**-0.33** (4.62)
ΔINST	**-6.65** (3.53)	**-6.42** (3.47)	**-6.44** (3.36)	**-5.89** (3.12)
SIZE	-3.41E-08 (0.36)		-1.13E-08 (1.18)	
RELSIZE	**-0.33** (3.88)	**-0.39** (6.05)	**-0.33** (3.80)	**-0.39** (6.48)
BM	0.004 (0.47)		0.005 (0.64)	
Log ILLIQ	0.083 (1.34)		0.093 (1.48)	
RUNUP	-0.33 (1.09)		**-0.65** (2.17)	**-0.67** (2.35)
HCINDX	-1.39 (1.50)		-0.82 (0.88)	
USINDX	1.05 (1.06)		0.98 (0.97)	
CAN	**1.06** (2.05)	**1.30** (3.00)	-0.56 (1.36)	
CORRUPT	-0.073 (0.64)		-0.040 (0.34)	
SOX	-0.068 (0.16)		-0.035 (0.08)	
Industry Fixed Effects	Yes	Yes	Yes	Yes
Adj. R^2	0.33	0.33	0.31	0.31

Table XII

Determinants of Foreign Cross-listing Excess Returns

Table XII reports regression results that examine the determinants of post-listing excess returns for foreign firms listing in the U.S. The dependent variable is the cumulative excess daily U.S. return computed between 70 and 250 trading days following cross-listing. The independent variables are as follows: *INST* is the percentage of shares outstanding held by U.S. institutional investors; *ΔINST* is the change in U.S. ownership between the listing quarter and four quarters after the listing date; *Residual INST* is the residual institutional ownership, calculated from (3c); *SIZE* is market capitalization one month prior to cross-listing; *RELSIZE* is *SIZE* divided by total market capitalization on the home exchange; *BM* is measured as book value in the year prior to cross-listing divided by *SIZE*; *DIV* is equal to 1 if the firm pays dividends in the year before cross-listing and 0 otherwise; *CORRUPT* is the 2007 Corruption Perception Index as calculated by Transparency International; *PE* is the ratio of the price one month prior to cross-listing divided by earnings in the year before cross-listing; *RUNUP* is the compounded daily excess return on the domestic exchange over the year prior to cross-listing; *HCINCX* and *USINDX* are the compounded daily returns in the home country and the U.S. over the year preceding cross-listing; *CAN* is a dummy variable equal to 1 if the cross-listing originates from Canada and 0 otherwise; *SOX* is a dummy variable equal to 1 if the stock cross-lists in 2002 or later. Industry fixed effects are defined on the basis of two digit SIC code. T-statistics are reported in parentheses and coefficients significant at the 10% level of significance appear in bold face. Model 1 (Model 3) presents the full specification using *INST* (*Residual INST*) and Model 2 (Model 4) retain variables from Model 1 (Model 3) significant at the 10% level of significance.

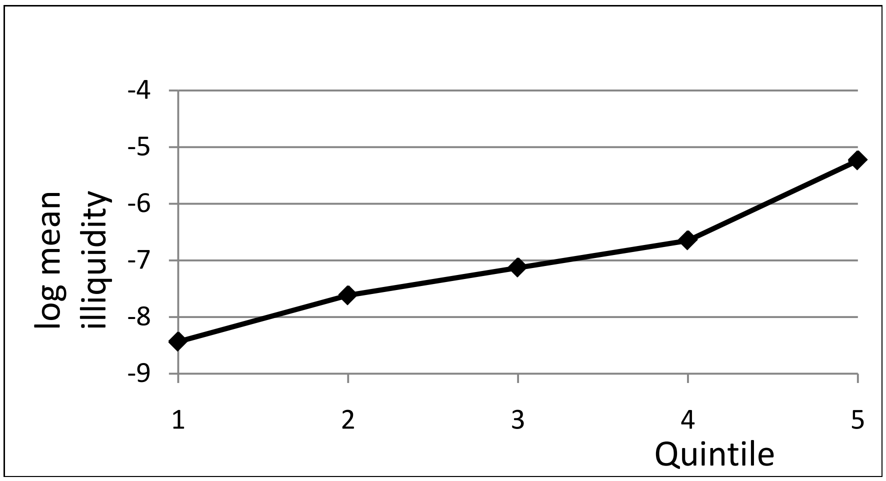
Independent Variable	Excess Return			
	Model 1	***Model 2***	***Model 3***	***Model 4***
Intercept	-0.17 (0.21)	-0.15 (0.61)	-0.53 (0.66)	**-0.72** (2.41)
Log INST	0.047 (1.52)	**0.054** (2.44)		
Residual INST			0.042 (1.22)	0.046 (1.37)
ΔINST	**2.54** (2.84)	**2.60** (2.96)	**2.49** (2.78)	**2.40** (2.69)
SIZE	-3.06E-11 (0.01)		1.06E-9 (0.24)	
RELSIZE	-0.056 (1.37)		-0.056 (1.36)	**-0.053** (1.85)
BM	0.005 (1.43)		0.0050 (1.37)	
Log ILLIQ	-0.0069 (0.23)		-0.008 (0.29)	
RUNUP	**0.36** (2.56)	**0.33** (2.46)	**0.41** (2.93)	**0.38** (2.81)
HCINDX	**-1.24** (2.83)	**-1.01** (3.35)	**-1.32** (3.03)	**-1.07** (3.25)
USINDX	0.22 (0.49)		0.24 (0.51)	
CAN	0.047 (0.19)		0.27 (1.41)	
CORRUPT	-0.070 (1.29)		-0.076 (1.39)	
SOX	-0.067 (0.34)		-0.071 (0.36)	
Industry Fixed Effects	Yes	Yes	Yes	Yes
Adj. R^2	0.14	0.14	0.13	0.13

Figure 3

Dispersion in Post Cross-listing Amihud Illiquidity Ratio and Cumulative Excess Returns

Panel A displays the Amihud Illiquidity Ratio and Panel B displays the annual cumulative excess returns after the cross-listing date for five equally populated bins ranked by these variables. The Amihud Illiquidity Ratio (Amihud, 2002) is defined as the mean of the absolute value of the weekly return scaled by the dollar value of weekly trading volume over 70 to 250 trading days after cross-listing. The cumulative excess return is calculated over the same window.

Panel A: Post-listing Amihud Illiquidity Ratio



Panel B: Post-listing Annual Cumulative Excess Returns

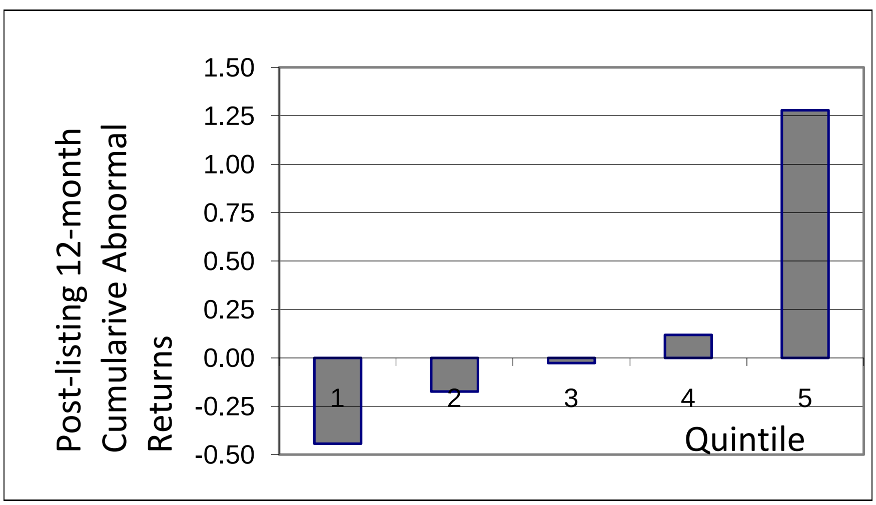
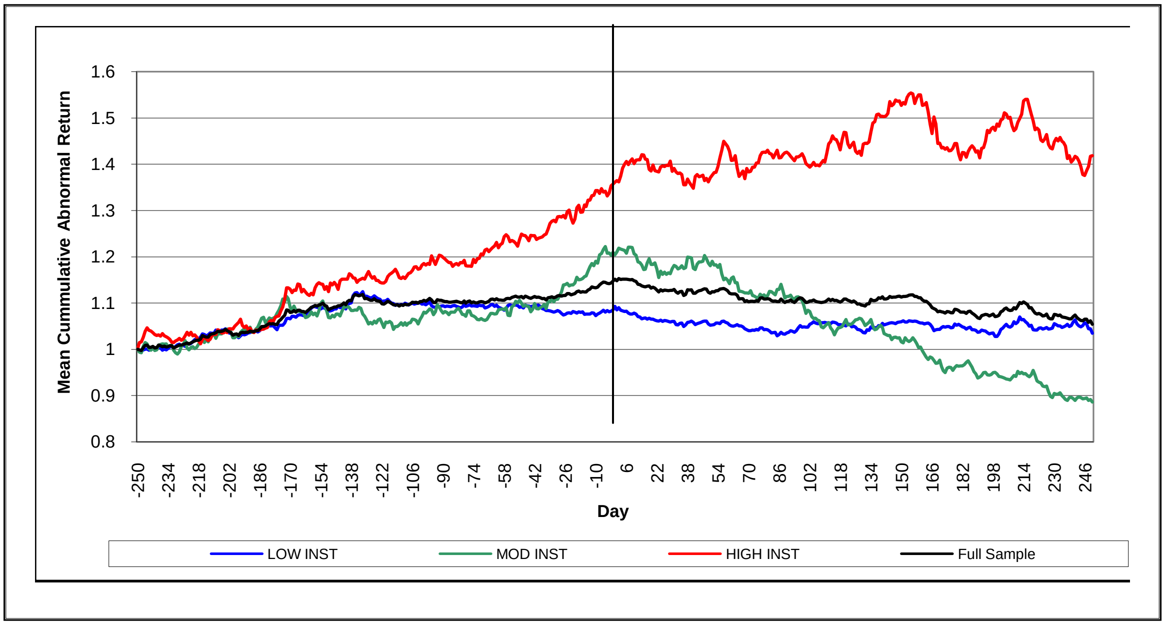


Figure 4

Cumulative Excess Returns Surrounding Cross-listing and U.S. Institutional Ownership

Figure 2 displays cumulative excess returns to terciles sorted by U.S. institutional ownership and for the full cross-listing sample. Cumulative excess return is calculated as the compounded value of the daily market-adjusted return starting one year prior to cross-listing, established as day -250 (the cross-listing is designated as day 0).



Chapter 4

Economic Conditions, Flight to Quality and Mutual Fund Flow

The notion of flight to quality has received substantial attention in the financial press. For instance, on November 12, 2007, Reuters News reports that:

> *"Investors put a net $20.5 billion into safe-haven money market accounts in the first week of November, reversing a brief flirtation with risk at the end of last month."*

The idea is that investors move to safer investments when economic conditions are expected to deteriorate and to riskier investments when economic conditions are expected to improve. Despite the intuitive appeal of flight to quality, there is little evidence of its prevalence or economic importance.[53] In this paper, we attempt to fill this void by examining the monthly asset allocation decisions of Canadian mutual fund investors. Our tests focus on the variation in aggregate flow to mutual funds that differ in terms of their riskiness as economic conditions change.[54]

On balance, the extant literature concludes that fund-level investments are largely sentiment driven rather than associated with economic fundamentals. Warther (1995), among many others, shows that investors base fund purchases on recent performance, chasing fund returns. This return chasing behavior is asymmetric in that investors fail to direct funds away from recent losers (see, for instance, Sirri and Tufano, 1998; Lynch and Musto, 2003). Additionally, funds that advertise recent success or receive greater media exposure attract a disproportionate share of the total flow (Sirri and Tufano, 1998; Jain and Wu, 2000).

[53] A recent exception is Beber et al. (2007) who examine flow in the bond market in times of market stress.

[54] While this issue could be examined in any context, our study of Canadian mutual fund flow provides some advantages. The number of fund categories is relatively small and clear-cut, allowing us to draw sharp inferences. Additionally, since the literature mainly studies U.S. mutual fund flow, our analysis provides evidence from a new arena.

In contrast to these fund-level studies, we examine how broad asset allocation changes in response to varying economic conditions. Specifically, as opposed to studying the relation between fund flow and fund performance, we investigate the relation between aggregate category flow and proxies for economic conditions. To capture economic conditions, we follow Fama and French (1989) and use TERM (the difference between the yields on long-term and short-term Canadian government bonds) and DEF (the difference between the yields on medium term corporate bonds and government bonds). To these, we add the short term interest rate (T-BILL, the 3 month Treasury Bill rate), as suggested by Shanken (1990). Economic conditions are good or expected to improve when TERM and T-BILL are high, or DEF is low, with the opposite holding for poor conditions.

Starting with monthly flow data for individual Canadian mutual funds over the period January 1991 through October 2005, we compute aggregate flow for seven major fund categories: Money Market; Bond; Balanced; Dividend and Income; Canadian Equity, U.S. Equity; and Foreign Equity. These categories account individually for at least 10% of overall monthly flow, on average, although each series shows considerable time-series variation. From the aggregate flow series, we calculate the percentage of overall flow for each category in each month. Our main tests then relate the percent of overall flow for the seven categories to the three proxies for economic conditions. By using percent flow, we are able to account for changes in the size of the market and exogenous shocks, and to make direct comparisons across fund categories.

Our results suggest that asset allocation decisions are influenced by economic conditions. When the Canadian economy is performing or expected to perform favorably (TERM is high, DEF is low, T-BILL is high), investors direct flow away from fixed income-type funds and towards equity based funds. For example, a one standard deviation increase in the term spread (1.13%) results in an 84% increase in percent flow to Canadian equity funds and a 74% decrease in flow to money market funds, relative to the previous month. The coefficients on DEF and T-BILL paint a broadly similar picture: taken together, Canadian investors allocate more (less) of their portfolios to risky than safe assets when economic prospects are good (poor). Moreover, the sensitivity to TERM, DEF and T-BILL is, for the most part, lower for balanced or dividend/income funds than for equity funds; thus, flow to the riskiest assets is most sensitive to economic conditions. This finding on the relation between aggregate flow and economic conditions is, to our knowledge, new.

Whether such a state-dependent asset allocation model enhances investor utility is ultimately determined by the risk-return profile that investors are able to achieve in relation to other strategies. Assuming a simple linear utility function, we consider the case of an investor who switches out of equity funds and into money market funds when TERM is in the lower quartile of its distribution and switches back into equity funds when TERM rises above this value.[55] We contrast this case with buy and hold strategies where the investor allocates wealth entirely to either equity or money market funds throughout the sample period. Our conclusions are

[55] A few brief comments are in order; Section III.b has the details. We choose a simple mean-variance utility function and vary the investor's risk aversion parameter. The investor is presumed to have a bias toward equity investing; hence, our choice of threshold for a move into or out of equities (TERM above or below the first quartile). Note that the economy is expected to do well when TERM is high and poorly when TERM is low.

sensitive to the weighting scheme used to calculate the return to the equity and money market asset classes. On an equal weighted basis, the buy and hold equity portfolio outperforms the flight to quality portfolio by 8.2% on a risk adjusted basis over the 14.75 year sample period. However, weighting by fund net asset value the relative performance of the two portfolios is reversed, as the flight to quality portfolio outperforms the buy and hold strategy by 5.4% on a risk adjusted basis. Based on either weighing methodology, the flight to quality portfolio realizes lower return volatility making it a more preferable strategy as investor risk aversion increases. It should be noted that this return comparison assumes consistent transaction costs between the equity and flight to quality portfolios. In reality, it is likely that the flight to quality portfolio would incur transaction costs in excess of the equity portfolio, reducing the relative performance of the flight to quality portfolio accordingly.

This evidence of flight to quality is reinforced by an analysis of three major episodes that occur over our sample period: the failure of Long Term Capital Management in August 1998; the Y2K problem in late 1999; and the terrorist attacks in September 2001. Even though these events are U.S.-centered, each episode is accompanied by a perceptible shift in Canadian mutual fund flow, away from riskier equity funds and into safer money market funds.

At the same time, a couple of results provide some pause. Controlling for economic conditions, we find persistence in flow at the broad asset level. Additionally, flow into equity funds tends to be higher following both negative and positive Canadian market returns, symptomatic of return chasing behavior. These results suggest that some part of aggregate level flow might have behavioral origins, as opposed to being rooted in macroeconomic analysis.

The remainder of this paper is organized as follows. Section 1 summarizes the relevant literature. Section 2 outlines the data and variables. Section 3 describes the results. Section 4 concludes.

I. Related Literature

Empirical research investigating the determinants of mutual fund flow can be divided into two broad groups. The first investigates the determinants of flow at the individual fund level. Several papers have documented a positive relation between fund flow and past performance (see, for example, Gruber (1996), Sirri and Tufano (1998), Patro (2006) and Cashman et. al. (2006a)). Further, Sirri and Tufano (1998) and Jain and Wu (2000) document that funds which advertise their success receive a disproportionate share of the inflow going to strong performers. The weight of the evidence indicates that the flow-performance relation is asymmetric. Sirri and Tufano (1998), Gruber (1996) and Lynch and Musto (2003) document that while investors send a larger proportion of flow to funds with strong performance they do not pull flow away from poorly performing funds. However, Cashman et al. (2006b) provide evidence that fund investors reduce flow to poor performers with the same intensity that they increase flow to strong performers.

Gruber (1996) and Zheng (1999) examine future returns to fund investors resulting from such performance chasing and find that funds with large inflow outperform funds with outflow. They interpret these results as suggesting that mutual fund investors have selection ability. Frazzini and Lamont (2007) provide contradictory evidence, documenting that the positive relation between fund performance and inflow is confined to short horizons of about one quarter. Over a longer window, funds with recent inflow realize significantly lower returns than those with outflow.

Finally, Cederburg (2008) examines return-chasing and subsequent performance for fund investors over the business cycle. He finds that, during expansions, investors earn higher risk-adjusted returns through return-chasing but investors do not chase returns during recessions, instead seeking funds with low market and book-to-market exposures.

Second, a more sparse literature examines the determinants of mutual fund flow at the aggregate level across fund types. Similar to the fund level research, Edwards and Zhang (1998) and Santini and Aber (1998) document that flow into equity funds is positively related to stock market performance. Santini and Aber (1998) also show that new money flow is negatively related to the lagged long term interest rate and positively related to contemporaneous personal disposable income. Campenhout (2004) finds that the change in the long rate, market return and fund performance are significant determinants of aggregate mutual fund flow in 11 European countries. Goetzmann et al. (1999) document that flow into equity funds is negatively correlated with flow into money market and precious metals funds. They argue this negative correlation suggests that fund allocations are not simply due to liquidity concerns but also reflect sentiment about the equity premium.

We contribute to this literature by examining the relation between aggregate mutual fund flow and economic conditions. In particular, we study the variation in relative flow to fund categories with different risk profiles as economic conditions change. By examining time-series variation in aggregate flow across fund classes, we are able to assess whether the risk preferences of mutual fund investors change with economic conditions. There is substantial evidence that aggregate mutual fund flows affects prices of the assets held by the funds and thereby also affect the performance of the funds themselves.[56] Thus, shifts in flow disrupt manager investment strategies, e..g. investor redemptions necessitate "fire sales" of the most liquid assets in the fund. For example, Coval and Stafford (2007) find that mutual funds in the top and bottom deciles of fund flows (high inflows or outflows to net assets) significantly underperform funds in the middle deciles. They further find that an investment strategy which short sells stocks likely to be involved in fire sales and which buys ahead of anticipated forced purchases earns an average annual abnormal return in excess of 15%. Based on this evidence, the performance of mutual funds is likely dependent on the ability of mutual fund managers to anticipate and prepare for economic cycle driven fund flow variations.

II. Sample and Variable Description

Mutual fund flow data are provided by the Investments Funds Institute of Canada (IFIC) which collects monthly sales, asset value and redemptions by fund for all Canadian mutual funds. Our dataset covers the period January 1991 through

[56] For instance, Braverman et al. (2007) and Warther (1995) document a positive relation between mutual fund flow and asset returns.

October 2005. Over our sample period, the number of funds expands from 430 to 1727. Funds tracked by the IFIC database are classified based on the 2007 Retail Investment Fund Category Definitions laid out by the Canadian Investment Funds Standards Committee (CIFSC). We exclude minor asset categories and focus on categories that account for 5% or more of total assets under management (aggregated across all funds) over the sample period. Application of this filter yields seven major asset categories: Canadian Equity (CE); Balanced (BA); Dividend and Income (DI); Bond (BO); Money Market (MM); U.S. Equity (US); and International Equity (IE). Appendix A provides the definitions of the seven fund categories. On average, the seven categories represent 92% of total assets under management over the 1991-2005 sample period.

Our objective is to examine variations in aggregate flow for these fund categories as economic conditions change. To this end, we are interested not in the level of the flow for each fund category, but rather in flow relative to the flow for other fund categories with different risks. Accordingly, we study the monthly percent flow for each category, calculated as:

$$percent\_sales_{j,t} = \frac{\sum_{i=1}^{N_j} net\_sales_{i,j,t}}{abs(\sum_{j=1}^{N}\sum_{i=1}^{N_j} net\_sales_{i,j,t})} \tag{4.1}$$

Here, subscript i denotes a fund, subscript j denotes a fund category and t denotes time. N_j is the total number of funds in category j, and N is the total number of fund categories (N=7). Net sales is defined as gross sales – gross redemptions + switches / transfers in – switches / transfers out. Distributions that are reinvested are not included in net sales. Thus, for example, percent sales in January 1991 for the Canadian equity fund category is the sum of net sales for all Canadian equity funds divided by the sum of net sales for all funds in the sample in January 1991.

By studying the percent of aggregate sales directed at a particular category, we are able to abstract from the effects of exogenous shocks and better isolate asset allocation effects. For instance, if business conditions worsen and income levels drop, investors are likely to reduce their investment in all categories. The question we are interested in is whether they reduce their investments proportionately across categories or if they target certain categories for more pronounced cutbacks, e.g. equity funds. This question is more cleanly addressed using the percent sales variable.

We relate the percent sales variable to three proxies for economic conditions. The term spread (TERM), is the difference between the yield on a long-term Canadian government bond (maturity of 10 years or longer) and the 3 month Canadian Treasury Bill rate. The default spread (DEF) is the difference between the yield on a portfolio of medium term Canadian corporate bonds and the yield on the medium maturity (three to five year) Canadian government bond. The final proxy is the yield on the 3 month Treasury Bill rate (T-BILL). We also include the previous month's market (Toronto Stock Exchange) return to capture the effects on flow of stock market conditions.

Fama and French (1989) show that TERM and DEF track economic conditions. Specifically, TERM is wide near business cycle troughs, when conditions are expected to improve, and narrow near peaks, when conditions are expected to worsen. DEF is wide when business conditions are poor and narrow when conditions are favorable. Chen (1991) shows that DEF predicts GDP growth over the following two quarters while TERM predicts GDP growth over the following five quarters. When DEF is high, slow growth is predicted; when TERM is high, rapid growth is predicted. Merton (1973) and Shanken (1990) suggest that T-BILL, the short-term rate, is a natural candidate for a state variable that captures variations in investment opportunities. If flight to quality is an important driver of flow, we expect to see the following coefficient signs in time-series regressions of flow on the proxies for economic conditions:

Decreasing Risk →

	CE	DI	BA	BO	MM
TERM	+	-	NP	-	-
DEF	-	+	NP	+	+
T-BILL	+	-	NP	-	-

NP means no prediction.

The predictions are clearest at the two ends of the risk spectrum, that is, for equity funds as well for fixed-income funds such as money market and bond funds. When business conditions are expected to improve, investors should increase exposure to equity funds and reduce exposure to fixed income funds. Investors should do the opposite when conditions are expected to deteriorate. TERM captures expectations regarding future economic conditions. Thus, for example, when TERM is high, investors are expected to overweight equity funds relative to fixed-income funds, so as to increase exposure during the pending economic expansion. Likewise, when TERM is low, investors should underweight equity funds and overweight fixed income funds. Accordingly, we expect a positive coefficient on TERM for Canadian equity and a negative coefficient for bond or money market funds.

DEF is a measure of prevailing business conditions, and is high when business conditions are poor and low when conditions are strong. Investors should reduce exposure to equity funds and seek a safe haven in fixed-income funds when DEF is high (i.e. conditions are poor), and increase allocations to equity funds and reduce allocations to fixed income funds when DEF is high. Thus, we expect to see a negative (positive) relation between Canadian equity funds (money market / bond funds) and DEF. The short term rate, T-BILL, is expected to be high when economic conditions are strong and low when conditions are weak. Therefore, predictions regarding the effects of T-BILL on flow are similar to those of DEF, with the signs reversed. In other words, the relation between Canadian equity funds and T-BILL is expected to be positive and the relation between money market or bond funds and T-Bill is expected to be negative.

Predictions related to balanced and dividend income funds are less clear. As balanced funds represent a blend of fixed income and equity funds, the relation between balanced fund flow and economic conditions is indeterminate. To the extent that dividends are sticky, dividend and income fund performance will be less sensitive to business cycle fluctuations. This argument suggests that dividend income funds might resemble fixed-income funds in terms of flow sensitivity to TERM, DEF and T-BILL.

We also examine flow to two categories of foreign investments, U.S. and foreign equity where foreign equity includes funds with greater than 95% equity holdings in companies outside Canada or the U.S. Considered relative to the domestic fund categories it is clear that foreign equity funds are of greater risk than domestic money market, bond or dividend income, but the risk ranking relative to Canadian equity funds is less clear. Given the high integration of the Canadian and U.S. economies and the high level of trade dependence of Canada on the U.S., TERM and DEF are highly correlated between the two countries. Given the high level of economic integration, TERM , DEF and TBILL would be expected to have similar relations to U.S. equity funds as predicted for Canadian equity funds.

Predictions regarding foreign equity funds are less clear. To the extent that all equity fund categories are riskier than money market or bond funds, we would expect a positive relation with TERM and TBILL and a negative relation with DEF for foreign equity funds. On the other hand, to the extent that economic conditions vary across international regions, times of poor economic performance in Canada may encourage investors to look abroad for investment opportunities, resulting in the opposite prediction.

The bond data are collected from two sources. Monthly data for Canadian government bonds and treasury bills are obtained from the Statistics Canada database. Data for the yields on medium term Canadian corporate bonds data are obtained from a database created by the Economist intelligence unit. These are investment grade corporate bonds, predominantly A to AA grade bonds, with maturity below 10 years. Data on the TSX market return are collected from DataStream.

III. Results

a. Economic conditions and aggregate mutual fund flow

We start by reporting descriptive statistics on flow for the seven mutual fund categories studied in this paper. Panel A of Table XIII reports statistics on the level of flow, assets and percent flow by fund category. Over the sample period, January 1991 to October 2005, the mean monthly dollar flow ranges from $190 million for the least active fund type (US funds) to $545 million for the most active fund type, international equity funds (all figures in Canadian dollars). Based on overall net assets, the largest category is Canadian equity funds (a mean of $60 billion), followed by international equities, balanced and money market funds. The medians yield generally similar orderings, although the medians are less extreme.

Figure 5 shows the mean aggregate monthly flow for each fund category by year. Net assets and flow grow rapidly over the sample period. For example, net assets under management in Canadian equity funds increase to 13 times their initial value, from $9 to $120 billion between 1991 and 2005. Money market and bond funds see similarly swift expansions, increasing five-fold and 11-fold respectively.

Panel B of Table XIII reports the mean number of funds in each category by year. The number of funds available to Canadian investors sees a large increase between 1991 and 2005. International equity funds have the greatest representation in 2005 (557) followed by Canadian equity funds (355) and balanced funds (275). Despite having, on average, 13% of total assets under management, there are only 87 money market funds in 2005. This indicates that money market assets are concentrated in larger funds relative to international or Canadian equity funds.

At the end of 2005, total assets under management in all Canadian funds amount to approximately $550 billion. In comparison, total assets under management in U.S. mutual funds are approximately $9 trillion (US) (ICI, 2007). Thus, the size of the Canadian mutual fund industry is approximately 1/20th the size of the US fund industry. Based on 2006 GDP, the U.S. economy is 13 times the size of the Canadian economy (IMF, 2007). Therefore, the sizes of the mutual fund industries in Canada and the US are roughly in line with the sizes of the two economies.

Table XIII also reports statistics on the percent sales variable, which is the key measure in our analysis. Over our sample period, the median percent sales values are, for the most part, tightly clustered, ranging from a low of 11% (for money market funds) to a high of 15% (for balanced funds), with the share of bond funds being 12% and those of Canadian equity and international equity funds being 14%. In this regard, the two smaller categories are U.S. equity funds and dividend and income funds, with median shares below 5%. The first quartile of percent sales is negative for two important categories (money market and Canadian equity funds), meaning that there is net outflow in approximately 44 of the 177 months in our sample. The inter-quartile range, (Q3-Q1)/2, is large for each category, e.g. 0.145 for

Canadian equity funds, and the standard deviation is even larger. In other words, there is considerable time-series variability in the relative flow series; it is this variation that we aim to explain using proxies for economic conditions.

Panel C of Table XIII provides the time-series correlation matrix for percent sales for the seven categories. Consistent with the risk ordering of the series, Canadian equity flow is positively correlated with international equity flow (0.51) and with balanced flow (0.49), and is negatively correlated with bond flow (-0.42) and money market flow (-0.57). Thus, investors appear to put money into, or pull money out of, equity funds at the same time. Moreover, when investors increase their allocation to equities, they reduce their allocation to fixed income (bond or money market) funds. This negative correlation between percent sales for Canadian equity funds and bond or money market funds is consistent with flight-to-quality effects.

The small negative correlation of -0.10 (not significant at conventional levels) between Canadian equity and dividend income suggests that the two categories are viewed as risk substitutes, possibly because dividend income is stable compared to capital gains. Flow to balanced funds is strongly negatively correlated with money market flow (-0.86) and positively correlated with dividend/income (0.65) and international equity (0.46) flow. To the extent that balanced funds are blends of stock and bond positions, the correlations are reasonable. Other notable correlations in flow exist for dividend and bond funds (0.67), dividend and money market funds (-0.47), bond and international equity funds (-0.53) and money market and international equity funds (-0.72). Taken together, these correlations provide preliminary evidence that flight to quality considerations are an important driver of investor allocations across fund categories with differing risk profiles. Shortly, we will carry out tests to formally examine the importance of flight-to-quality effects.

First, Table XIV provides summary statistics on the independent variables in these tests. Descriptive statistics are shown in Panel A. TERM, the yield premium for investing in long-term over short-term bonds, averages approximately 2% per year. The mean value of DEF, the premium for investing in risky relative to safe bonds, is 1.2% per year. These values are similar to the mean U.S. values reported by Fama and French (1989) for TERM (1.99%) and DEF (0.96%) over a longer sample period. T-BILL, the annualized T-Bill rate, averages 4.6% between 1991 and 2005. Finally, the mean TSX return is 70 basis points per month. There is appreciable variability in each series, seen in the large standard deviation or inter-quartile range.

Panel B reports the time-series correlations among these variables. DEF and TERM have a correlation of 0.36, and the fact that the correlation is well below 1.0 implies that they capture different aspects of economic conditions (as argued by Fama and French (1989) and Chen (1991)). The correlation of -0.54 between TERM and T-BILL is due to the presence of the 3 month interest rate in both series.[57] DEF and T-BILL are negatively correlated, although the coefficient is relatively small (-

[57] For robustness we replicate our results with an alternative specification of TERM using 1-3 year Canadian Government Bond data in place of the 3 month Canadian Government Bond to mitigate potential co-linearity biases between TERM and T-BILL and our results and conclusions are unaffected.

0.10). Last, the correlations between the TSX return and the contemporaneous values of TERM, DEF and T-BILL are small and insignificantly different from zero.

We now turn to the central issue in the paper, the importance of flight-to-quality effects in driving asset allocation decisions. To that end, Table XV presents the results of regressions of PERCENT SALES, the percent sales for the seven fund categories, on DEF, TERM, and T-BILL (each measured at the end of the previous month), as well as the lagged value of PERCENT SALES and the previous month's TSX return.[58] DEF, TERM and T-BILL are proxies for economic conditions. If flight-to-quality effects are important, higher values of TERM and T-Bill or lower values of DEF (reflecting improved economic condition) will be associated with an increasing share of flow to riskier categories. The previous month's percent sales for the category in question is included to capture persistence in aggregate flow. The TSX return from the previous month is included to examine the possibility of return chasing at the aggregate level. This should lead to a positive (negative) relation between the lagged market return and equity (bond) flow. We separate positive and negative market returns to look for asymmetric return chasing effects.

We start with the flow to Canadian equity funds (CE). CE is positively related to TERM and T-BILL, and negatively related to DEF. The coefficients on TERM and T-BILL are significant at better than the 1% level of significance while that on DEF is significant at the 10% level. Since an increase in TERM or in T-BILL, or a decline in DEF, signifies improvements in economic conditions, this means that the share of aggregate flow directed at equity funds increases when economic conditions

[58] We scale the dependent and independent variables by the standard deviation of each time series. This standardization allows us to directly assess economic significance.

improve. The coefficients similarly imply that equity fund flow declines when economic conditions deteriorate, as seen in declining TERM or T-BILL or increasing DEF. This relation between aggregate equity fund flow and the proxies for economic conditions is consistent with the flight-to- quality story.

At the other extreme, aggregate percent sales for Canadian money market funds, MM, is negatively related to TERM (at the 10% level of significance), and positively related to DEF (p-value < 0.01). The coefficient on T-BILL is insignificantly different from zero. These coefficients imply that the MM share of flow increases when economic conditions are expected to deteriorate and declines when they are expected to improve. This, too, is consistent with flight to quality.

The risk levels of dividend/income (DI), balanced (BA) and bond (BO) funds lie somewhere between the extremes represented by CE and MM. If flight to quality considerations are important, we expect investors to increase their allocation to bond funds and perhaps reduce that to dividend/income and balanced funds as conditions become gloomier, and do the opposite as conditions improve. For BO, the coefficients on TERM and T-BILL are negative and the coefficient on DEF is positive. Thus, relative flow into bond funds increases when economic conditions are expected to worsen (i.e. as TERM or T-BILL declines or DEF increases). For BA, the coefficient on TERM is significantly above zero at the 10% level of significance, but this coefficient is only half as large as that for equity funds. Investors appear to regard balanced funds as broadly similar to stock funds; however, their bond holdings lead to a more muted reaction to changing economic conditions. For DI, the only significant coefficient is that on T-BILL. The fact that this coefficient is negative suggests that dividend/income funds are viewed more as bond than as stock funds. Again, however, the coefficient is half as large as that for equity flow.

The predictions regarding international equities (IE) and US equities (US) are less clear. We expect the dominant effect to be associated with risk aversion on the part of Canadian investors. In other words, flow to all risky investments, including non-Canadian funds, should decline as economic conditions get worse. On the other hand, if economic conditions in Canada and the US or international markets do not overlap completely, investors might increase their allocations to US or international funds in search of (relative) safety when Canadian economic conditions worsen. Looking at the coefficients for these two categories, it appears that flow for both IE and US reacts similarly to that for purely Canadian equity funds. That is, the coefficients on TERM and T-BILL are positive (for US flow, the coefficient on T-BILL is insignificantly different from zero), while the coefficient on DEF is negative. Thus, the share of mutual fund investments going to international and U.S. equities declines as Canadian economic conditions worsen. This is consistent with increased risk aversion on the part of Canadian investors resulting in larger flow to safe investments compared to any risky investment.

Turning to the control variables, we see that the coefficient on lagged percent sales is significantly above zero. Consequently, there is persistence in the share of aggregate flow going to each category. This is consistent with evidence of persistence in flow at the fund level (see, for example, Warther, 1995). The effects of the market return variables are significant only for CE. The coefficients imply that investors send a larger fraction of their flow to Canadian equity funds both after the market rises and after it falls, and are consistent with return chasing behavior at the aggregate level.

b. Does flight to quality pay?

The regression evidence indicates that investors' asset allocation decisions reflect a flight to quality motivation. Specifically, they direct a greater proportion of their dollars toward risky investments (Canadian equity, US and international equity funds) and a smaller fraction to safer investments (bond and money market funds) in the face of improving economic conditions in Canada, and do the opposite when conditions are expected to worsen. In this section, we consider whether this behavior benefits investors

By varying wealth allocations across broad asset classes as a function of economic conditions, investors may realize benefits stemming from higher returns or lower risk. If investors expect equity returns to be, on average, low (or negative) when economic conditions are poor, shifting wealth from equity to fixed income investments in anticipation of an economic downturn may improve portfolio returns. Alternatively, since stock market volatility tends to increase during downturns (Schwert, 1989), investors may be able to reduce portfolio risk by investing in fixed income assets in advance of downturns.

To analyze the risk – return implications of flight to quality behavior, we construct three portfolios. The first two portfolios, which serve as benchmarks, see the investor allocating 100% of his wealth to Canadian equity funds and Canadian money market funds respectively (CE and MM portfolios, respectively). If the investor holds all his wealth in Canadian equity (money market) funds, we assume that the investor will receive the return to the median Canadian equity (money market) fund in each month.[59] The third portfolio, which results from flight to quality behavior (FTQ portfolio), normally is 100% invested in Canadian equity funds. However, the investor shifts all of his wealth into money market funds when TERM drops below its first quartile value, and shifts back into equity funds when TERM rises above its first quartile value. We condition on TERM because it is the most significant predictor of flow in Table XV.[60] The distribution of TERM is recalculated in 'real time', using the five most recent years of monthly bond market data.[61] This makes the analysis more realistic by basing the allocation decision on information actually available to the investor. We choose the first quartile as the cutoff value for TERM because we assume that an investor has a bias toward investing in equities.

This asset allocation rule results in a 100% allocation to the CE portfolio for much of the January 1991 - October 2005 sample period, excepting the three periods shown in Figure 6. During these three periods, wealth is allocated entirely

[59] We calculate monthly fund returns as: return =net asset/(lag net assets + net sales) -1 as recommended by the IFIC. This calculation assumes flows to each fund are realized at the start of each period when in reality in/out flows are realized throughout the month. As we are unable to determine the inter-,monthly timing of flows, we are not able to calculate a time weighted monthly return for each fund which would more accurately approximate actual fund returns. Based on the assumption that fund flows follow a random walk process, under or over estimation of fund returns is equally likely and thus our results are not systematically biased. As an additional precaution we utilize median returns to mitigate the potential for extreme returns resulting from our return estimation process to bias our results. Results are similar and our conclusions are unaffected if mean returns are used instead.

[60] We obtain similar results and our conclusions are unaffected if we instead condition on either DEF or TBILL.

[61] In other words, at the end of each month, we drop the most distant month of data, and add the most recent month of data, for the long-term bond and short-term yields. Then, we compare the most recent value of TERM to the first quartile of the updated distribution of TERM.

to the MM portfolio. Table XVI reports the holding period return, mean monthly return and standard deviation of monthly return for the three portfolios, where the portfolio returns are calculated over the full sample period. Table XVI also presents the Sharpe Ratio for the three portfolios, calculated as the ratio of the mean return less the risk free rate to the standard deviation of the return.[62]

We start by comparing the FTQ and CE portfolios. Our conclusions depend to a certain extent on the weighing factor utilized when calculating the monthly portfolio return. On an equal-weighted basis, the CE portfolio realizes a higher holding period return (257% vs. 171%) and a statistically significant higher average return (1.45% vs. 0.97%) than the FTQ portfolio.[63] Thus, one dollar invested in January 1991 would grow to $258 in October 2005 if invested in the CE portfolio, compared to $172 if the investor switches between equity and money market funds according to TERM (FTQ portfolio). The monthly standard deviation for the FTQ portfolio (2.98) is lower than that for the CE portfolio (3.93). However, this differential is not as pronounced as the difference in mean returns. Thus, the performance differential present in the Sharpe Ratio is also present when the mean excess return is standardized by portfolio standard deviation. In this scenario, the FTQ strategy underperforms a passive buy-and-hold equity investment strategy.

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[62] As a proxy for the risk free rate we use the mean risk free rate for the U.S. over the sample period as reported on Ken French's website. Given the degree of integration between the U.S. and Canadian economies we feel this is an appropriate proxy. The closest Canadian equivalent would be the yield to 3 month Canadian Government bonds, but the return to this instrument as a risk free proxy is likely overstated due to its use by Canadian banks for liquidity management.

[63] The t-statistic from a two sample t-test comparing the mean monthly return to the Equity and Flight to Quality portfolios is 2.64.

By contrast, on a value weighted basis, the performance differential between the CE and FTQ portfolios is reduced.[64] The average monthly return for the CE and FTQ portfolios (1.04% and 0.91%, respectively) are now statistically indistinguishable. Given the lower standard deviation of monthly returns for the FTQ portfolio (2.96 vs 3.74), the Sharpe Ratio is 0.24 for the CE portfolio compared to 0.26 for the FTQ portfolio. Thus, the FTQ portfolio now marginally outperforms the CE portfolio on a risk adjusted basis. It should be noted that the return and Sharpe Ratio comparison are completed excluding transaction costs, which would likely be higher for the FTQ portfolio. If transaction costs were factored into the analysis, the margin of superior performance of the FTQ portfolio would likely be reduced or potentially the CE portfolio would regain superiority. We believe that value weighting more realistically captures the likely return an investor would realize via the FTQ strategy as this methodology most heavily weights the most widely held funds in each category.

Table XVI also summarizes the return statistics if the investor holds only money market funds for the entire sample period (MM portfolio). On an equal-weighted basis, the mean monthly return of 0.4% and holding period return of 71% are considerably lower than those for the CE and FTQ portfolios. This lower return comes with a benefit, in the form of a low return standard deviation of 2.4% per month. However, the Sharp Ratio is 0.12, much lower than the ratios for the CE and FTQ portfolios. When we weight fund returns by net asset value, the performance differential between the MM portfolio and the CE and FTQ portfolios is reduced, but

[64] Fund net asset value is used as the value weighting factor.

this comes at the cost of higher risk. Regardless of the weighing scheme, the switching strategy is superior to a buy and hold money market fund strategy.

We can gain additional perspective on the risk-reward tradeoff by using utility analysis. We assume a simple mean-variance form for investor utility:

$$U_i = R_i - \frac{1}{2} A \sigma_i^2 \tag{4.2}$$

where U_i is investor utility, R_i is the mean monthly return for portfolio *i* and σ_i is the standard deviation of monthly return for portfolio *i*, and *A* is the coefficient of absolute risk aversion.

We allow the risk-aversion parameter, *A*, to vary between 0 and 20. Figure 7 displays the variation in investor utility across the three portfolios for different levels of *A*. The broad conclusions are similar for the equally-weighted and value weighted cases. For an investor with low risk aversion, the equity buy-and-hold strategy yields higher utility. For risk aversion levels in excess of 15 (5) on an equal weighed (net asset value weighted) basis, the FTQ portfolio provides higher utility. In relation to the MM portfolio, the investor realizes higher utility from the FTQ portfolio for all risk aversion levels considered. Note that, for the MM investment strategy, utility is negative for risk aversion levels in excess of 13.5 (11.5) in the case of the equal weighted (net asset weighted) portfolio. While investors with high enough risk aversion (e.g. greater than 22 on an equal weighted basis) will derive higher utility from the MM portfolio than the FTQ portfolio, the fact that utility is negative means that they would shun both classes of investments.

c. Major shocks and flight to quality

In closing, we take a slightly different look at flight to quality by considering three major shocks that occur over our sample period: the failure of the hedge fund Long-Term Capital Management in August 1998; the Y2K crisis in late 1999; and the terrorist attacks on New York City on Sept 11, 2001. Each of these events was accompanied by fears of a meltdown in global financial markets. If these events triggered flight to quality concerns, we might see their effects in Canadian mutual fund flow. We examine funds at the extremes of the risk spectrum, Canadian equity and money market funds.

Figure 8 shows the percent flow to the CE and MM portfolios between June 1997 and March 2002, a time-frame which includes the three crises. The results are striking. In each case, we see a strong increase in the percent flow to the MM and a drop in the flow to the CE portfolio. In August 1998, the month of the LTCM debacle, the percent flow to the MM portfolio is 2.43, while that to CE portfolio is -0.80, indicating a net inflow of $1,850 million for money market funds and an outflow of $627 million for Canadian equity funds. Over the four months preceding Y2K, there is a mean monthly percent inflow for the MM portfolio of 0.98 and this reverses in January 2000. The opposite pattern is apparent for the CE portfolio, with a mean percent outflow over the same four-month period of 2.20. The equivalent dollar flow in the four months preceding Y2K is $1.6 billion into money market funds and

$1.4 billion out of Canadian equity funds. Last, in September and October 2001, the months including and following the 9/11 attacks, the MM portfolio experiences percent inflows of 1.16 and 0.94 and the CE portfolio sees percent outflows of 0.13 and 0.0001. These translate into cumulative two-month inflow of $500 million for money market funds and outflow of $300 million for Canadian equity funds.

Note that, with the possible exception of Y2K, the episodes are U.S. centered. However, each episode was accompanied by fears that financial markets all over the world would face difficulties. Our evidence suggests that Canadian mutual fund investors were mindful of these risks and transferred money from risky to more secure investments. Analysis of these three episodes confirms that flight to quality considerations are an important driver of investor asset allocations.

IV. Concluding Comments

In this paper, we examine the asset allocation decisions of mutual fund investors. We are interested in the importance of flight to quality considerations as a driver of fund flow, i.e. whether investors direct money towards safer (riskier) investments when economic conditions are expected to become weaker (stronger). With this goal in mind, we study monthly net flow for the universe of Canadian mutual funds between 1991 and 2005. We separate funds into seven categories—Canadian Equity, Dividend and Income, Balanced, Bond, Money Market, U.S. Equity and International Equity—and aggregate the flow for each of these categories. Our variable of interest is the percent flow for each of the seven categories. At extreme ends of the risk spectrum are Canadian Equity and Money Market funds, with Balanced, Dividend and Income, and Bonds representing intermediate risks.

As proxies for economic conditions, we use the default spread (DEF), term spread (TERM) and short term interest rate (T-BILL). Following prior research (e.g. Fama and French, 1989; Chen, 1991), we assume that economic conditions are good when DEF is low, and TERM and T-BILL are high. We then relate the percent flow for the seven categories to DEF, TERM and T-BILL, plus controls. Our main finding is that an expected improvement in Canadian economic conditions causes investors to direct flow away from fixed income funds and towards equity funds; when conditions are expected to deteriorate, the reverse happens. For example, a one standard deviation increase in the term spread (1.13%) results in an 84% increase in the percent of aggregate flow to Canadian equity funds, and a 74% decrease in the percent flow to money market funds, relative to the previous month. Based on net sales in October 2005, these changes translate into an extra monthly inflow of $84 million for all Canadian equity funds and an outflow of $74 million for Canadian money market funds.

Whether this flight to quality asset allocation rule is beneficial or detrimental to mutual fund investors can be determined by examining the risk-return profile that they are able to achieve in relation to other strategies. Accordingly, we compare the case of an investor who switches out of equity funds and into money market funds when TERM is in the lower quartile of its distribution and switches back into equity funds when TERM rises above this value. We contrast this case with the case where he stays either in equities or in money market funds for the whole sample period.

Our conclusions depend to a certain extent on the weighing methodology utilized. On an equal weighted basis returns to the flight to quality portfolio are on average 0.48% lower per month than a simple buy-and-hold equity fund portfolio. This average monthly return difference culminates into an 86% higher return to the equity buy-and-hold portfolio across our sample period of just less than 15 years. If instead, monthly returns are weighted by net asset value, which more accurately reflects the returns realized by funds more heavily favoured by investors, the difference in returns to the portfolio is statistically insignificant. Regardless of the weighting methodology, return volatility is lower for the flight to quality portfolio making it increasingly preferable with increasing investor risk aversion.

Last, we examine the flow to equity and money market funds surrounding three major crises: the Long Term Capital Management debacle, the Y2K problem and the 9/11 terrorist attacks. These three episodes prompted fears of a meltdown in global financial markets: if they triggered flight to quality concerns, we might see their effects in Canadian mutual fund flow. We find that each episode is accompanied by significant flow into money market funds and out of equity funds. For example, the August 1998 Long Term Capital Management failure sees Canadian investors move $1,850 million into money market funds and $627 million out of equity funds.

The message from our analysis is that mutual fund investors appear to take into consideration the information contained in signals of the economy's health (captured by variables such as DEF, TERM and T-BILL) while making their asset allocation decisions. Specifically, investors direct their dollars at asset categories on the basis of the risk characteristics of the category in conjunction with the prevailing economic environment. This state-dependent asset allocation model does not unambiguously dominate a buy and hold equity investment strategy over our sample period: our conclusions are sensitive to the choice of asset weighting scheme and risk aversion parameter. However, investors end up with a superior return-risk tradeoff and with higher utility relative to passive investment in money market funds.

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Appendix A

The fund type classifications utilized in this chapter are established based on the Canadian Investment Funds Standards Committee (CIFSC) 2007 Retail Investment Fund Category Definitions (CIFSC, 2007). Following is the definition of each of the fund types examined in this chapter.

Balanced Funds (BA)
Funds in the Balanced Funds group must invest between 5% and 90% of their non-cash assets invested in Equity Securities and between 10% and 95% of their non-cash assets in fixed-income securities.

Canadian Fixed Income (BO)
Funds in the Canadian Fixed Income category must invest at least 90% of their fixed income holdings in Canadian dollars with an average duration greater than 3.5 years and less than 9.0 years. In addition, these funds must invest primarily in investment-grade fixed-income securities, such that the average credit quality of the portfolio as a whole is investment grade (BBB or equivalent rating or higher) and not more than 25% of the portfolio's holdings are invested in high yield fixed income securities. For purposes of the category definition, up to 30% of a Fund's assets may be held in Foreign Fixed Income products which will be treated as Canadian content provided that the currency exposure on those holdings is hedged into Canadian Dollars.

Canadian Equity (CE)
Funds in the Canadian Equity category must invest at least 90% of their equity holdings in securities domiciled in Canada, and their average market capitalization must be greater than the Canadian small/mid cap threshold.

Canadian Dividend & Income Equity (DI)
Funds in the Canadian Dividend & Income Equity category must have a stated mandate to invest primarily in income-generating securities and must invest at least 90% of their equity holdings in securities domiciled in Canada. In addition, these funds must invest at least 50% of their non-cash assets in income-generating securities such that the 3-year weighted average yield on the equity component of the fund's portfolio is at least 1.5 times the average yield of the Canadian Equity Fund benchmark, defined as the S&P/TSX Equity Index. The fund's average capitalization must exceed the Canadian small/mid cap threshold.

International Equity (IE)
Funds in the International Equity category must invest at least 95% of their equity assets in countries other than Canada and the United States and at least 70% of their equity assets in developed countries.

Money Market Funds (MM)
Funds in the Money Market group must invest at least 95% of their total net assets in cash or cash equivalent securities and otherwise comply with the legal definition of Money Market funds as outlined in National Instrument 81-102.

U.S. Equity (US)
Funds in the U.S. Equity category must invest at least 90% of their equity holdings in securities domiciled in the United States, and their average market capitalization must be greater than the U.S. small/mid cap threshold.

Table XIII

Panel A of Table XIII reports descriptive statistics for monthly net sales, net assets and percent sales. Each variable is monthly in frequency reported at months end from January 1991 through November 2005. Net sales is aggregate net sales for all funds in each category defined as gross sales – gross redemptions + switches / transfers in – switches / transfers out. Distributions that are re-invested are not included in net sales. Net assets is aggregate net assets for all funds in each category defined as the value of all holdings of the fund less all liabilities. Percent sales is calculated as net sales for each fund type divided by the absolute value of the sum of net sales to all funds. Panel B reports the mean number of funds in each fund type by year. Panel C reports the correlation matrix of percent sales across fund types, values significant at conventional levels (α=0.05) appear in bold face.

Panel A: Descriptive Statistics of Fund Variables by Fund Type

Variable	Type	Mean	Median	Q1	Q3	STD
Net Sales ($x10^6$)	BA	516	287	72	726	712
	BO	345	317	73	586	383
	IE	545	388	183	794	1028
	CE	437	267	-41	799	874
	DI	284	113	22	451	448
	MM	347	342	-323	873	869
	US	191	101	25	297	299
Net Assets ($x10^9$)	BA	45	53	18	65	30
	BO	24	25	12	31	14
	IE	60	67	27	89	37
	CE	66	81	29	96	36
	DI	18	18	4	23	16
	MM	33	32	16	48	16
	US	17	17	4	31	13
Percent Sales	BA	0.23	0.15	0.07	0.26	0.62
	BO	0.19	0.12	0.05	0.22	0.44
	IE	0.26	0.14	0.02	0.30	1.35
	CE	0.07	0.14	-0.03	0.26	0.56
	DI	0.22	0.04	0.01	0.17	0.52
	MM	-0.18	0.11	-0.14	0.42	2.18
	US	0.09	0.05	0.01	0.10	0.33

Panel B: Summary of number of funds per fund type, by year

BA	Canadian Balanced Funds	IE	International Equity Funds
BO	Canadian Fixed Income Funds	MM	Money Market Funds
CE	Canadian Equity Funds	US	United States Equity Funds
DI	Canadian Dividend and Fixed Income Funds		

**Refer to Appendix A for detailed fund type definitions

Year	BA	BO	IE	CE	DI	MM	US
1991	61	65	51	112	13	54	30
1992	67	74	67	132	15	66	43
1993	74	84	84	144	18	71	54
1994	82	95	134	165	21	72	63
1995	96	110	208	190	32	78	76
1996	93	114	239	198	33	77	81
1997	102	112	255	210	33	80	87
1998	119	119	277	245	37	84	98
1999	135	124	310	274	42	86	115
2000	160	131	464	306	44	87	154
2001	174	132	681	327	43	90	193
2002	197	138	717	347	49	91	234
2003	214	137	675	362	65	87	283
2004	252	143	616	355	74	87	277
2005	275	147	557	355	79	87	257

Panel C: Correlation Matrix of Percent Sales by Fund Type

	CE	BA	DI	BO	MM	IE	US
CE	1						
BA	**0.491**	1					
DI	-0.101	**0.650**	1				
BO	**-0.422**	**0.175**	**0.667**	1			
MM	**-0.570**	**-0.863**	**-0.470**	0.039	1		
IE	**0.507**	**0.457**	-0.140	**-0.531**	**-0.722**	1	
US	0.009	**-0.165**	-0.093	-0.139	-0.078	0.018	1

Table XIV

Panel A of Table XIV reports descriptive statistics for the independent variables used as proxies for the economic state in Canada. The correlation matrix of the same variables is included in Panel B, values significant at conventional levels (α=0.05) appear in bold face. The variables are monthly in frequency reported at months end from January 1991 through November 2005. DEF is the difference in yield between medium term corporate bonds and 3 to 5 year Canadian Government Bonds. TERM is the difference in yield between the 10 year plus Canadian Government Bond and the 3 month Canadian Treasury Bill. T-BILL is the yield on the 3 month Canadian Treasury Bill. TSX RETURN is the average daily return to the TSX Composite Index over the month of interest.

Panel A: Descriptive Statistics of Independent Variables

Variable	N	Mean	Median	Q1	Q3	STD
DEF	177	1.21	1.13	0.87	1.52	0.40
TERM	177	2.09	2.22	1.15	3.03	1.13
T-BILL	177	4.64	4.63	2.86	5.62	1.90
TSX RETURN	177	0.007	0.01	-0.01	0.04	0.04

Panel B: Correlation Matrix of Independent Variables

	TERM	DEF	T-BILL	TSX RETURN
TERM	1			
DEF	**0.360**	1		
T-BILL	**-0.541**	**-0.097**	1	
TSX RETURN	0.022	-0.010	-0.010	1

Table XV

Percent Sales Time-series Regression Results

Table XV reports time-series regression results of monthly percent sales to each fund type regressed on lagged percent sales, proxies for economic state and measures of stock market performance. The dependent variable is percent sales by fund type. Percent sales is calculated as net sales for each fund type divided by the absolute value of the sum of net sales to all funds, where net sales is aggregate net sales for all funds in each category defined as gross sales – gross redemptions + switches / transfers in – switches / transfers out. Distributions that are re-invested are not included in net sales. TERM is the difference in yield between the 30 year plus Canadian Government Bond and the 31 day Canadian Treasury Bill at month end. DEF is the difference in yield between medium term corporate bonds and 3 to 5 year Canadian Government Bonds at month end. T-BILL is the yield on the 31 day Canadian Treasury Bill at month end. NEG TSX RETURN is equal to TSX RETURN if TSX RETURN is <0 and is otherwise equal to zero, where TSX RETURN is the average daily return in each month to the TSX Composite Index. POS TSX RETURN is calculated in the corresponding fashion but in relation to a positive value for TSX RETURN. VOLAT is the mean daily square return in each month. Standardized coefficient values are report with t-statistics reported in brackets below each coefficient. Coefficients significant at conventional levels (α=0.10) appear in bold face.

BA	Canadian Balanced Funds	IE	International Equity Funds
BO	Canadian Fixed Income Funds	MM	Money Market Funds
CE	Canadian Equity Funds	US	United States Equity Funds
DI	Canadian Dividend and Fixed Income Funds		

**Refer to Appendix A for detailed fund type definitions

Fund Level Percent Sales Regression Results

Independent Variables	Dependent Variable: Percent Sales by Fund Type						
	CE	DI	BA	BO	MM	IE	US
INTERCEPT	**-1.168**	0.362	0.459	0.536	0.104	-0.144	0.624
	(3.40)	(0.85)	(1.65)	(1.82)	(0.23)	(0.29)	(0.92)
LAG TERM	**0.324**	0.153	-0.009	-0.128	**-0.230**	0.224	0.235
	(4.06)	(1.55)	(0.16)	(1.95)	(2.13)	(1.93)	(1.48)
LAG DEF	-0.0943	-0.058	0.050	**0.136**	0.118	**-0.222**	**-0.297**
	(1.48)	(0.70)	(1.10)	(2.48)	(1.29)	(2.23)	(2.22)
LAG T-BILL	**0.289**	-0.076	**-0.154**	**-0.195**	-0.031	0.207	0.078
	(3.83)	(0.81)	(2.65)	(2.98)	(0.30)	(1.82)	(0.50)
LAG NEG TSX RETURN	**-0.222**	0.005	0.076	0.047	0.026	-0.032	-0.00
	(3.07)	(0.07)	(1.22)	(0.73)	(0.34)	(0.47)	(0.00)
LAG POS TSX RETURN	0.135	-0.022	-0.049	-0.059	0.013	0.045	0.036
	(1.78)	(0.29)	(0.72)	(0.86)	(0.18)	(0.70)	(0.49)
LAG % SALES	**0.415**	**0.299**	**0.669**	**0.615**	**0.231**	**0.227**	**-0.389**
	(6.13)	(4.12)	(11.82)	(10.09)	(3.12)	(3.02)	(5.56)
R^2	0.27	0.21	0.27	0.38	0.19	0.32	0.02

Table XVI

Portfolio Risk – Return

Table XVI reports the risk and return for three portfolios; the equity and money market portfolios allocate 100% of wealth to equity and money market funds respectively, and realize the median monthly return to each respective fund type. The median monthly return to the two asset types are determined using equal weighting in Panel A and net asset value weighting in Panel B. The Flight to Quality portfolio allocates 100% of wealth to equity funds but transfers wealth to money market funds when TERM drops below is first quartile value based on the previous five years of monthly bond data. TERM is the difference in yield between the 30 year plus Canadian Government Bond and the 3 month Canadian Treasury Bill at month end. Holding period return is the return to each portfolio from January 1991 to October 2005. Mean Return is the mean monthly return and Standard Deviation is the standard deviation of monthly returns for each portfolio over the same timeframe. The Sharpe Ratio is mean return less the risk free rate divided by standard deviation, where the risk free rate is the monthly average risk free rate over the sample period (0.125%).

Panel A: Equal Weighted

	Money Market	**Equity**	**Flight to Quality**
Holding Period Return (%)	71.36	257.49	171.15
Mean Return (%)	0.40	1.45	0.97
Standard Deviation (%)	2.41	3.93	2.98
Sharpe Ratio	0.12	0.34	0.28

Panel B: Net Asset Value Weighted

	Money Market	Equity	Flight to Quality
Holding Period Return (%)	127.18	184.02	161.90
Mean Return (%)	0.72	1.04	0.91
Standard Deviation (%)	3.51	3.74	2.96
Sharpe Ratio	0.17	0.24	0.26

Figure 5

Mean Monthly Aggregate Net Assets by Fund Type and Year

Figure 5 reports the trend in mean monthly aggregate net assets for all funds in each category from 1991 – 2005. The averages calculated for 2005 exclude data for November and December due to data availability constraints. Net assets is defined as the value of all holdings of the fund less all liabilities.

BA	Canadian Balance Funds	IE	International Equity Funds
BO	Canadian Fixed Income Funds	MM	Money Market Funds
CE	Canadian Equity Funds	US	United States Equity Funds
DI	Canadian Dividend and Fixed Income Funds		

**Refer to Appendix A for detailed fund type definitions

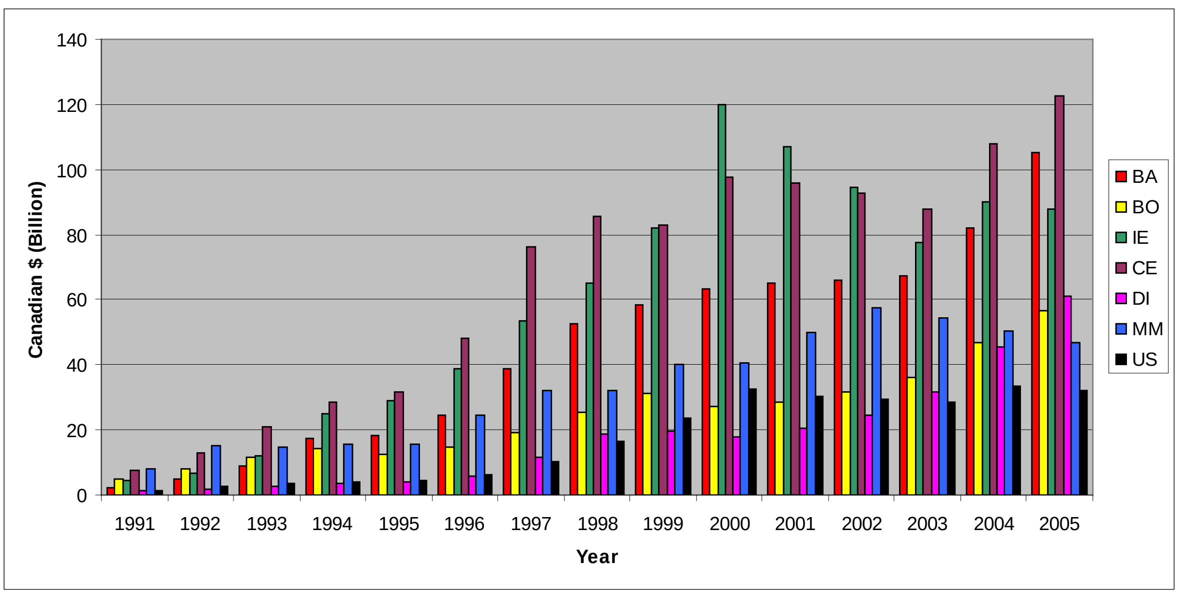


Figure 6

Transition Periods of the Flight to Quality Portfolio

Figure 6 plots the TSX Composite Index overlaid with transition periods of the flight to quality index. The shaded area indicate periods when the investor would place 100% of wealth into money market funds, the remainder of time 100% of wealth is invested in equity funds.

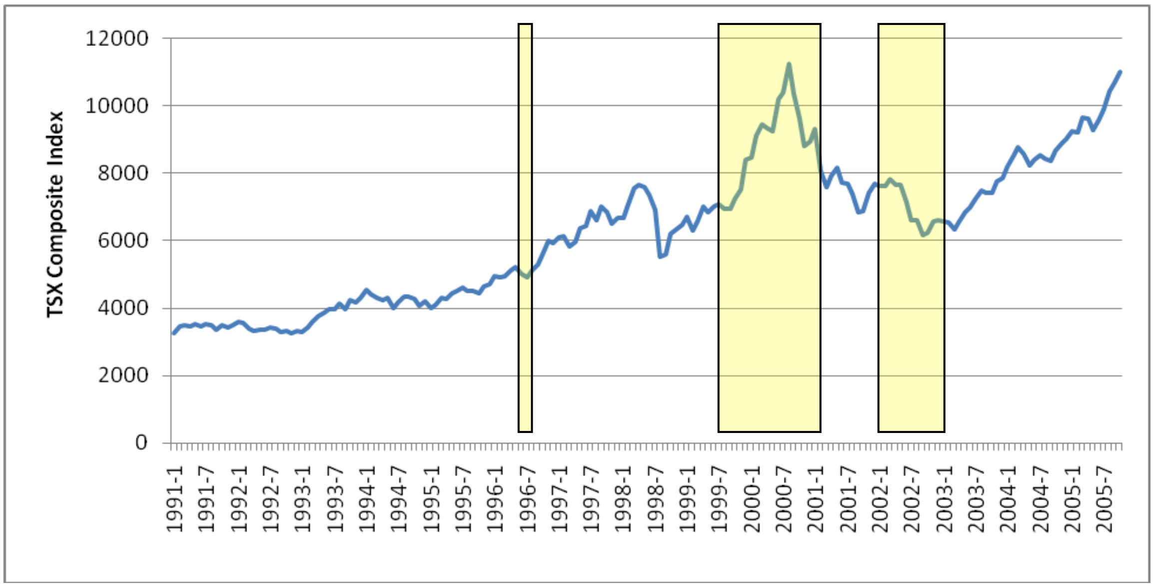
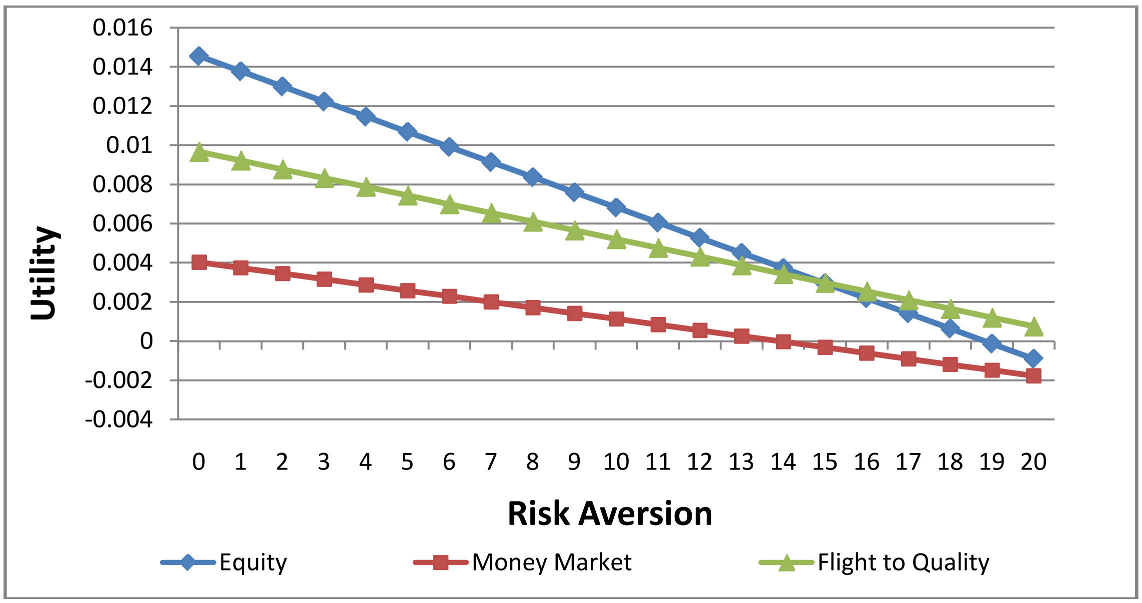


Figure 7

Variation of Investor Utility across Portfolios

Figure 7 displays the variation of investor utility with variation in the coefficient of risk aversion across three portfolios; the equity and money market portfolios allocate 100% of wealth to equity and money market funds respectively and realize the mean monthly return to each respective fund type. The Flight to Quality portfolio allocates 100% of wealth to equity funds but transfers wealth to money market funds when TERM drops below is first quartile value based on the previous 5 years of monthly bond data. TERM is the difference in yield between the 30 year plus Canadian Government Bond and the three month Canadian Treasury Bill at month end.

Panel A: Equal Weighted



Panel B: Net Asset Value Weighted

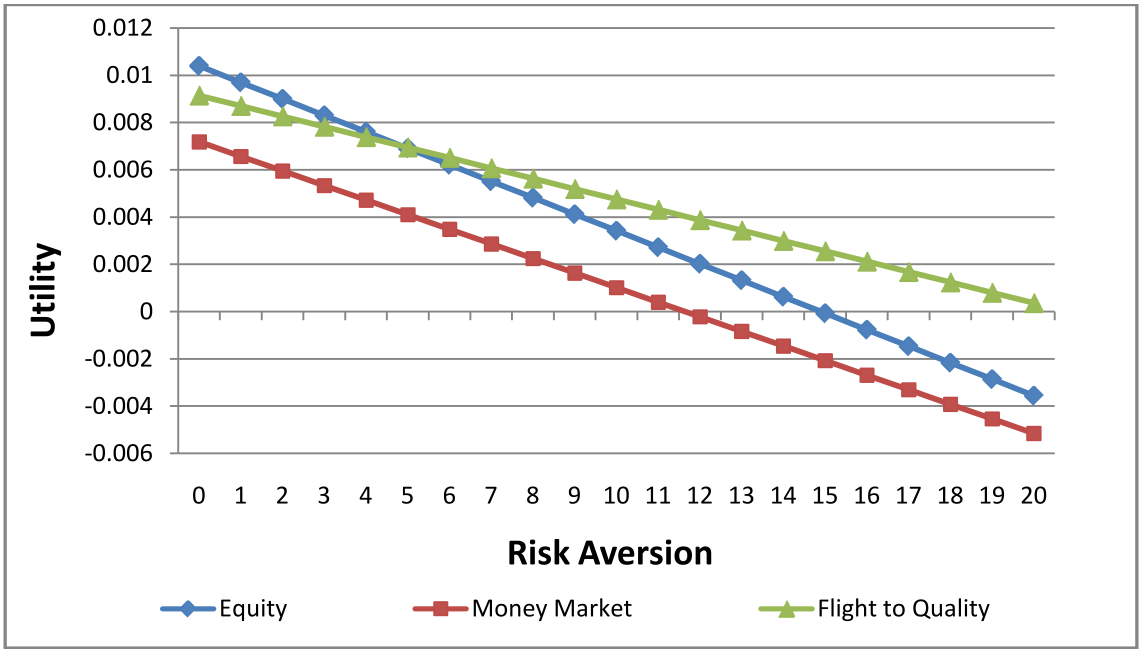


Figure 8

Percent Sales for Canadian Equity and Money Market Funds During Crises

Figure 8 displays monthly percent sales to Canadian Equity (CE) and Money Market (MM) fund types from June 1997 to March 2002 over which time there were three significant global events which influenced financial markets: 1) August 1998, the hedge fund Long Term Capital Management (LTCM) lost 44% of total assets becoming a prominent example of the risk potential in the hedge fund industry. 2) Third and fourth quarters 1999, fears surrounding the potential effect of the turn of the century on the date tracking systems in computers (Y2K). 3) September 11, 2001, terrorist attacks on the World Trade Center in New York City (9/11). Percent sales is calculated as net sales for each fund type divided by the absolute value of the sum of net sales to all funds, where net sales is aggregate net sales for all funds in each category defined as gross sales – gross redemptions + switches / transfers in – switches / transfers out. Distributions that are re-invested are not included in net sales.