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The Timing and Terms of Management Buyouts for State-owned Enterprises in China

Zhuming Chen[1]

*Department of Finance and Investment, School of Business,
Sun Yat-sen University, Guangzhou, China, 510275

K. C. Chen

Department of Finance and Business Law, The Craig School of business,
California State University, Fresno, CA 93740-8001, U.S.A.

Abstract

The reform of state-owned enterprises (SOEs) has been at the heart of all economic transformations in China during the last decades. The recent policy position of the central government toward state-owned enterprise reform is privatization and diversification of ownership. Although property rights transfer can take on many forms such as mergers and acquisitions, we focus only on management buyouts in this paper. Using a real options setting, this paper theoretically examines the timing and terms of management buyouts for SOEs in China.

JEL: G13, G34.
Keywords: state-owned enterprise; management buyout; China; real option; property rights transfer; MBO; SOE

I. Introduction

Since 1980, China has launched a number of schemes to reform its economic systems[2]. The reform measures were designed to create a modern enterprise system based on the Company Law, which defines a state-owned enterprise's (SOE hereafter) property right to dispose of its asset, including the equity of the state and the state's limited responsibility to its shareholders in proportion to its capital contribution. According to Lin and Zhu (2000), the central goal of the shareholding reform is to establish a "modern enterprise system" featuring corporate governance structures that separate the government from enterprises. The separation is deemed necessary for SOEs to achieve full autonomy in structural and operational decisions. It is also hoped that, by quantifying equity ownership, ownership restructuring will help facilitate efficient reallocation of capital resources through mergers and acquisitions (M&A).

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Since the reform of state-owned enterprises has been at the heart of all economic transformations in China, the recent policy position of the central government toward SOE reform is privatization and diversification of ownership. As the central government has further transferred control rights over SOEs to local governments, the local governments are virtually entitled to implement whatever reform methods they think fit. It has been long recognized that the fastest way to transform poorly performing SOEs into a "modern corporate system" is to sell part or all of the assets or equity interest of the SOE to foreign investors or other strategic partners. According to Milman (1999) who studies M&A activities in China, two forms of transactions had dominated during 1985-1996: acquisition of majority interest and acquisition of partial interest. The relatively low volume of asset acquisitions and mergers in the region could be explained by the fact that stock acquisitions of majority or partial interest were legally less sophisticated than mergers and asset acquisitions, and could be expeditiously executed under proper intermediation.

However, the progress of SOEs since reform has been very uneven: some SOEs are quite modern in their organization and management, but others have failed to embrace the realities of a market economy. According to the statistics reported by the Chinese State-owned Assets Supervision and Administration Commission (SASAC), there were 150,000 SOEs in China in 2003, among which 98 percent were small and medium-sized SOEs[3]. However, during the decade-long reform process, state-assets stripping has taken place hand in hand with the adoption of various reform measures. Lee and Hahn (2004) document that the problem of insider control has led to asset stripping during the last decade. Xu (2000) estimates that approximately 30~100 billion yuan worth of state assets flew out of the state treasury every year. This figure would have been much higher if the loss of state assets incurred in the process of transferring property rights was included. According to Xu, inadequate asset evaluation has been one of the main sources of state assets stripping in the process of corporatization and forming and operating joint ventures.

The undervaluation scandals are more notorious in many of the management buyout (MBO) transactions, because in an MBO where agency problems could be particularly acute, management along with the aid of third-party investors has a motive to acquire the SOE at the lowest possible price. The unfairness revealed in the recent upsurge in MBOs can be attributed to share mispricing, illegal sources of funding, and lack of transparency. Lee and Hahn (2004) document several MBO transactions during 1999 and 2002, which are listed in Table 1, showing that many firms' transfer prices are lower than their net asset values. They also discover that in most cases, share transfer prices are settled without any revaluation of assets, resulting in the stripping of state property and infringement on minority shareholder rights.

In this paper, we will theoretically analyze the timing and terms of management buyouts for SOEs in China. Although property rights transfer can also take on many forms such as mergers and acquisitions, we only focus on MBOs for two reasons. First, in recent years, MBOs have increased not only in both size and number, but also across continents[4]. In China today, more and more local governments choose to privatize SOEs to insiders through MBOs and employee stock ownership plans (ESOPs). According to Lee and Hahn (2004), as of June 2003, about two hundred out of nine hundred listed firms in which the state held shares had plans to conduct MBOs. Second, considerable research has studied the timing and terms of mergers. For instance, Lambrecht (2004) analyzes mergers motivated by economies of scale using a real option setting with endogenous timing. He shows that merger synergies exhibit call option-like characteristics, and he derives closed form solutions for the merger timing and for the post-merger ownership shares, respectively. Morellec and Zhdanov (2004) incorporate competition and imperfect information into their model when determining the terms and timing of takeovers by solving

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[2] For literature review on economic reforms in China, see Dodd (1996), Lin and Germain (2003), Lin and Zhu (2000), and Xu (2000).
[3] See http://www.sasac.gov.cn

Table 1
Transfer prices and net asset value per share for eight MBOs during 1999-2002

Firms	**Management's ownership (%)**	**Transfer price**[1] ***(yuan)***	**Net asset value per share *(yuan)***	**Pricing rule**
Shengli	17.65	2.27 2.27	2.24 2.27	Net asset value
Tebian Diangong	27.64	1.24 2.5 3.1	3.1 3.28 3.28	Not announced
Fosu	29.48	2.95	3.3	Not announced
Yutong Bus Manufacturing	100	Not announced	7.0	Not announced
Fangda Group	36.1	3.08 3.28 3.55	3.43 3.43 3.45	Not announced
Guangdong Meide	30.59	Not announced Not announced Not announced 2.95 3.0	2.99 3.31 3.56 3.99 3.99	Not announced
Dongting Shuizhi	22.9	5.75	5.84	Net asset value
Ordos	43.8	1.77 1.77	5.64 5.79	Not announced

Source: Annual and mid-term reports (1999-2002) of listed companies in Lee and Hahn (2004).
[1] MBO was conducted at various stages.

option exercise games between bidding and target shareholders.

The rest of the paper is organized as follows. Section II presents the basic assumptions and the theoretical framework. In Section III, we will derive the optimal timing and terms of selling an SOE via management buyout and report the comparative statistics of the finding. Section IV concludes.

[4] See MergerStat Review (2000) in the U.S., Filatotchev, Starkey, and Wright (1994) in Central and Eastern Europe, and Baruch and Woodward (1998) in the U.K.

II. Assumptions and the Theoretical Framework

In the following analysis, a real options setting with endogenous timing is used following Dixit and Pindyck (1994) and Lanbrecht (2004). We first assume that the Chinese government will sell poorly performing SOEs via MBOs and retain those that are efficiently managed. Following Fradkin (2004) who studies the put option that a firm has on a real asset, the poorly performing SOE will be sold as soon as the present value of future cash flows equals its salvage value. That means the Chinese government in fact holds a put option with exercise price set below the SOE's current value. On the contrary, the manager involved in the MBO deal has a call option to buy the SOE at the lowest possible price. The manager is often rewarded with an in-the-money call option for two plausible reasons. First, it is not uncommon in China where agency problems are notoriously prevalent in corporate governance. Second, it could be the local government's desperate intent to sell the SOE in compliance with the central government's economic reform policies.

Next, denote firm g as the pre-MBO SOE and firm m as the post-MBO SOE. The firm's equity value, $E_j(pt)$ (j=g,m), is determined by a stochastic output price, p_t. When engaging in the buyout, a sunk cost denoted by x_j (j=g,m) is incurred that refers to fees paid to investment bankers and lawyers.

The firm's equity value prior to MBO includes a put option for the intended seller and a call option for the manager as follows:

$$E_j(p_t) = V_j(p_t) + Q_j(p_t), \quad j = g, m \tag{1}$$

where $V_j(p_t)$ denotes the stand-alone equity value without the option and $Q_j(p_t)$ denotes the endogenously determined value of the option to sell or buy.

We also assume that the SOE pre- and post-MBO operates in the same business and employs a Cobb-Douglas production function, $L_i^a K_i^b$, with an instantaneous profit function given by:[5]

$$p_t L_j^a K_j^b - w_L L_j \tag{2}$$

where L_j represents labor cost with per unit cost denoted by ω_L and K_j denotes capital.

The SOE is a price taker with its output price, p_t, following a geometric Brownian motion:

$$dp_t = \mu p_t dt + \sigma p_t dB_t \tag{3}$$

where $\mu < r, \sigma > 0$, Bt is a standard Brownian motion, σ is the instantaneous volatility, and r is the risk–free interest rate.

We further assume that the SOE is equity-financed with a 100% payout ratio and the investors are risk neutral. Following Dixit and Pindyck (1994), the stand-alone value of the firm is simply the present value of all future profits discounted at the risk-free interest rate as follows:

$$V_g(p_t) = \frac{\Pi K_g^\theta p_t^\gamma}{r - \mu\gamma - \frac{1}{2}\sigma^2\gamma(\gamma-1)} = \frac{\Pi K_g^\theta p_t^\gamma}{r - f(\gamma)} \tag{4}$$

subject to $f(\gamma) < r$, where $\gamma = \frac{1}{1-a} > 1$, $\theta = \frac{b}{1-a}$, and $\Pi = (a^{\frac{a}{1-a}} - a^{\frac{1}{1-a}})\omega_L^{\frac{-a}{1-a}} > 0$.

[5] The Cobb-Douglas production function displays decreasing returns to scale with respect to the variable input (i.e., $0 < a < 1$) but increasing returns to scale when both inputs are variable (i.e., $a + b > 1$).

III. The Timing and Terms of an MBO

When analyzing the timing and terms of an MBO, it is equally important to recognize that each participating party owns an option linked to the deal and both participants will exercise their options only when reaching an agreement on the buyout price that maximizes their respective objective function. Therefore, the timing of buyout can be derived from a strategic equilibrium in which each party determines an optimal option-exercise strategy while taking into account the other party's exercise strategy.

Since the SOE holds a put option which is denoted as Q_g , following Dixit and Pindyck (1994) we can construct a risk-free portfolio that contains the option to sell (a put option) and a long position in firm's output. Applying standard real option valuation methods, it is shown below that there exists a threshold, $\bar{p}_g$, such that exercising the put option is optimal as soon as the state variable, p_t, falls below $\bar{p}_g$, and Q_g is the solution to a second order differential equation subject to a number of boundary conditions.

The value of the put option, Q_g, must satisfy the following equilibrium condition:

$$rQ_{gt} = \frac{d}{d\Delta} E_t Q_{g(t+\Delta)} \Big|_{\Delta=0} \tag{5}$$

Applying Ito's lemma to Q_{gt} yields

$$rQ_g(p_t) = \mu p_t Q_g^{'}(p_t) + \frac{\sigma^2}{2} p_t^2 Q_g^{''}(p_t) \tag{6}$$

The solution to the above differential equation is given by

$$Q_g(p_t) = A_1 p_t^{\lambda} + A_2 p_t^{\beta} \tag{7}$$

where λ and β, are positive and negative roots to the characteristic equation, respectively.

The value of the put option, $Q_g(p_t)$, is determined by the following three boundary conditions. First, the put option value goes to zero when $p_t \to \infty$. When the output prices rises, the SOE should perform well and the government will not sell it, leaving the put option worthless. So, $A_1 = 0$. Second, the value-matching condition at $\bar{p}_g$ requires that

$$Q_g(\bar{p}_g) = A_2 \bar{p}_g^{\beta} = I_g - V_g(\bar{p}_g) - X_g \tag{8}$$

where I_g is the selling price imposed by the SOE. Substituting equation (4) into equation (8) and rearranging yields

$$A_2 = (I_g - \frac{\Pi K_g^{\theta} \bar{p}_g^{\gamma}}{r - f(\gamma)} - X_g)(\frac{1}{\bar{p}_g})^{\beta} \tag{9}$$

Finally, applying the smooth-pasting condition to equation (8) at $\bar{p}_g$ obtains

$$\beta A_2 \bar{p}_g^{\beta-1} = -\frac{\Pi K_g^{\theta} \gamma \bar{p}_g^{\gamma-1}}{r - f(\gamma)} \tag{10}$$

Dividing equation (8) by equation (10) and rearranging, the optimal selling threshold for the SOE, $\bar{p}_g$, can be derived as follows:

$$\bar{p}_g = \left(\frac{\beta}{\beta-\gamma}\cdot\frac{(I_g - X_g)(r - f(\gamma))}{\Pi K_g^{\theta}}\right)^{\frac{1}{\gamma}} \tag{11}$$

In an MBO, the manager has a call option to buy the SOE. The manager's optimal exercise strategy can be derived by showing that there exists a threshold, $\bar{p}_m$, such that exercising the call option is optimal as soon as the state variable, p_t, exceeds $\bar{p}_m$, and the value of the call option, Q_m, is the solution to a second order differential equation similar to equation (6) as follows:

$$Q_m(p_t) = C_1 p_t^{\lambda} + C_2 p_t^{\beta} \tag{12}$$

The manager's call option value relies on the following conditions. First, the value of call option goes to zero as $p_t \rightarrow$. So, $C_2 = 0$. Second, because MBO per se could not generate any operational synergy, we assume that the manager possesses some special skills and expertise, thus making managerial synergy the sole incentive for the manager to pursue the MBO. We further assume the manager's ability can be measured by a constant α, which can be observed by the government and the public, so there is no asymmetric information.

Similar to equation (4), the value of the post-MBO firm is given by:

$$V_m(p_t) = \frac{\Pi(\alpha K_g)^{\theta} p_t^{\gamma}}{r - f(\gamma)} \tag{13}$$

The benefit that the manager can extract from the post-MBO firm is

$$V_m(p_t) - V_g(p_t) = \frac{\Pi((\alpha^{\theta} - 1)K_g^{\theta} p_t^{\gamma}}{r - f(\gamma)} \tag{14}$$

The benefit is positive if $\alpha > 1$. To justify why $\alpha > 1$, we can think of a firm-specific shock as the trigger of this relative ability enhancement. This shock could be induced by a regulatory change, e.g., China's economic reform to privatize SOEs as discussed in Section I, that unleashes the manager's inner potential coupled with managerial freedom without abiding by rigid government policies. Another supporting evidence for $\alpha > 1$ is provided by Kaplan (1989), who finds that MBOs' post-buyout operating improvements and value increases are due to "improved incentives" rather than wealth transfers from employees or inside information.

The value-matching condition at $\bar{p}_m$ requires that

$$Q_m(\bar{p}_m) = C_1 \bar{p}_m^{\lambda} = V_m(\bar{p}_m) - I_g - X_m \tag{15}$$

where Xm is the transaction cost incurred by the manager in the MBO. Substituting equation (13) into equation (15) and rearranging gets:

$$C_1 = \left(\frac{\Pi \alpha^{\theta} K_g^{\theta} \bar{p}_m^{\gamma}}{r - f(\gamma)} - (I_g + X_m)\right)\left(\frac{1}{\bar{p}_m}\right)^{\lambda} \tag{16}$$

Applying the smooth-pasting condition at $\bar{p}_m$, we obtain

$$\lambda C_1 \bar{p}_m^{\lambda-1} = \frac{\Pi \gamma \bar{p}_m^{\gamma-1} \alpha^{\theta} K_g^{\theta}}{r - f(\gamma)} \tag{17}$$

Dividing equation (15) by equation (17) and rearranging, the optimal purchase threshold for the MBO, $\bar{p}_m$, can be derived as follows:

$$\bar{p}_m\bar{p}_m = \left(\frac{\lambda}{\lambda-\gamma}\cdot\frac{(I_g + X_m)(r - f(\gamma))}{\Pi\alpha^{\theta}K_g^{\theta}}\right)^{\frac{1}{\gamma}} \quad (18)$$

Since in equilibrium the negotiated outcome of the MBO must satisfy the constraint, $\bar{p}_g = \bar{p}_m = \bar{p}_m^*$, by setting equation (11) equal to equation (18) we can derive the equilibrium MBO selling price, I_m^*:

$$I_m^* = \frac{\beta(\lambda-\gamma)\alpha^{\theta}X_g + \lambda(\beta-\gamma)X_m}{\beta(\lambda-\gamma)\alpha^{\theta} - \lambda(\beta-\gamma)}$$

$$= \frac{\frac{\beta}{\beta-\gamma}\cdot\alpha^{\theta}X_g + \frac{\lambda}{\lambda-\gamma}\cdot X_m}{\frac{\beta}{\beta-\gamma}\cdot\alpha^{\theta} - \frac{\lambda}{\lambda-\gamma}} \quad (19)$$

and the MBO trigger price, $\bar{p}_m^*$:

$$\bar{p}_m^* = \left(\frac{\beta\lambda}{\Pi K_g^{\theta}}\cdot\frac{(r - f(\gamma))(X_g + X_m)}{\beta(\lambda-\gamma)\alpha^{\theta} - \lambda(\beta-\gamma)}\right)^{\frac{1}{\gamma}}$$

$$= \left(\frac{\beta}{\beta-\gamma}\cdot\frac{\lambda}{\lambda-\gamma}\cdot\frac{r-f(\gamma)}{\Pi}\cdot\frac{X_g+X_m}{K_g^{\theta}}\cdot\frac{1}{\frac{\beta}{\beta-\gamma}\alpha^{\theta} - \frac{\lambda}{\lambda-\gamma}}\right)^{\frac{1}{\gamma}} \quad (20)$$

For the above results to be meaningful,

$$\frac{\beta}{\beta-\gamma}\cdot\alpha^{\theta} - \frac{\lambda}{\lambda-\gamma} > 0 \quad (21)$$

where $\frac{\beta}{\beta-\gamma}$ and $\frac{\lambda}{\lambda-\gamma}$ are the hysteresis factors of selling and buying, respectively. Because MBO generates no operating synergies, the left-hand side of equation (21) is the surplus gained from the managerial synergy induced by improved incentives.

It is worth noting that the comparative statistics with respect to the MBO threshold in equations (19) and (20) are consistent with Lambrecht's (2004) findings with respect to the merger threshold. Specifically, the equilibrium MBO selling price has no direct relation with the SOE's capital and is a decreasing function of the manager's ability. In other words, the greater the manager's ability is, the lower the selling price. Since in an MBO the manager has a strong motive to acquire the SOE at the lowest possible price to reduce out-of-pocket money investment, conflicts of interest are therefore inevitable.

In addition, higher transaction cost will not only raise the MBO selling price, but also delay the MBO transaction. Larger surplus $\frac{\beta}{\beta-\gamma}\cdot\alpha^{\theta} - \frac{\lambda}{\lambda-}$ which the manager can extract from the MBO, also expedites the MBO transaction. Similarly, increasing uncertainty (σ) will delay the MBO and an increase in drift μ of the output price will speed up the MBO. Increasing the parameters θ and γ via increasing the return to scale parameters a and b will speed up MBO. Finally, bigger capital also speeds up the MBO, which implies that bigger SOEs will be sold first, ceteris paribus.

V. Conclusion

In this paper, we theoretically analyze the timing and terms of management buyouts for SOEs in China. At the outset, we assume that the Chinese government possesses a put option to sell the poorly performing SOE as soon as its condition deteriorates, and that the manager of the SOE is given a call option to buy the firm at the lowest possible price. When analyzing the timing and terms of an MBO, it is important to recognize that each participating party owns an option linked to the deal and both participants will exercise their options only when reaching an agreement on the buyout price that maximizes their respective objective function.

Based on the work of Dixit and Pindyck (1994) and Lambrecht (2004), we use continuous-time real options techniques to model the dynamics of MBO. The timing and terms of a management buyout is derived from a strategic equilibrium in which each party determines an optimal option-exercise strategy while taking into account the other party's exercise strategy. In addition, the comparative statistics with respect to the MBO threshold are consistent with Lambrecht's (2004) findings with respect to the merger threshold.

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The Turn-of-the-Month Effect in Stocks Trading on the National Stock Exchange of India

Steven Freund
UMASS Lowell
College of Management
One University Avenue
Lowell, MA 01854
(978) 934-2818
steven_freund@uml.edu

Ravi Jain
UMASS Lowell
College of Management
One University Avenue
Lowell, MA 01854
(978) 934-2854
ravi_jain@uml.edu

Yash Puri*
UMASS Lowell
College of Management
One University Avenue
Lowell, MA 01854
(978) 934-2807
yash_puri@uml.edu

* Contact author

Abstract

We test for the turn-of-the-month calendar anomaly in stocks that trade on the National Stock Exchange of India, using the daily returns of the S&P CNX Nifty index for the period 1992 through 2004. The mean and the median daily returns are found to be significantly higher for trading days which occur during the turn-of-the-month period. Using regression models, we verify the presence of a turn-of-the-month effect, even after controlling for potential January and weekend effects.

JEL Classification Numbers: G11, G12, G14
Key words: Market Efficiency, Stock Market Anomalies, Turn-of-the-Month Effect

I. Introduction

Calendar anomalies in stock returns are studied in a large number of research papers. These regularities include time of the day, day of the week, time of the month, and month of the year. Other observed calendar anomalies include the turn of the month and turn of the year, and the days before or after a holiday, when the exchange is closed.

Calendar anomalies are in direct conflict with one of the major paradigms of financial economics, the efficient market hypothesis (EMH) of Fama (1965, 1970). Although the earliest papers on regularities appeared more than 70 years ago in Fields (1931) and Wachtel (1942), research on U.S. stock market anomalies accelerated after monthly and daily data became widely available from the Center for Research in Security Prices (CRSP).

The search for calendar anomalies in global equity markets followed shortly thereafter. Although the availability of CRSP data facilitates the research on anomalies in the U.S. markets, it also leads to warnings of potential selection bias and data mining from repeated use of the same data.[1] Lakonishok and Smidt (1988) mention data snooping as a potential explanation for calendar anomalies, and suggest that the use of new data sources is the best remedy for avoidance of this particular statistical error. One benefit of examining a new global market for calendar anomalies is that it is an out-of-sample test. As pointed out by Agrawal and Tandon (1994), every new data set which supports the same anomaly increases the likelihood that the finding is not related to data snooping.

In this paper, we focus on stocks trading on the National Stock Exchange of India. (NSE). In particular, we test for the turn-of-the-month (TOTM) effect, the finding that stock returns are greater as one month ends and another begins. Our data for this study is the daily returns of the S&P CNX Nifty, a stock index of stocks trading on the NSE. We find a significant TOTM effect in the Nifty index returns, even after we control for potential January and weekend effects.

The remainder of this paper is organized as follows: Section II consists of a literature review of calendar anomalies in general, with an emphasis on the U.S. and global evidence related to TOTM research; Section III provides a brief review of the S&P CNX Nifty on the National Stock Exchange of India; Section IV presents our methodology and findings and Section V concludes.

II. Literature Review

The earliest calendar anomaly considered during the post-EMH era of modern finance was the January effect.[2] First suggested in Wachtel (1942), the January effect is the tendency of stock returns in January to exceed the returns in the other months of the year. Wachtel provides two possible explanations. Stocks which have decreased in price may be sold in December causing the price to decrease even more. The subsequent rebound in January may be one explanation for the January effect. Second, an increased demand for cash during the holiday season in December may also artificially depress the prices of securities in December.

Rozeff and Kinney (1976) provide the first rigorous study of the January effect, and show that the average monthly return for January is three percent greater than the average return for the other eleven months of the year, using an equal-weighted index of New York Stock

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[1] See Lo and Mackinlay (1990) and Fama (1991).

[2] The variation of stock returns over various calendar periods is often referred to as a "seasonality". In Thaler (1987b) it is suggested that this is a misnomer since the time period under question is not always related to the "seasons". It is for this reason that we favor the term "calendar" anomaly, which would seem to be appropriate for all of these regularities except for ones associated with the time of the day.

Exchange stocks. The fact that Rozeff and Kenney use an equal-weighted index is of consideration, because it is shown in Banz (1981) that small firms outperform large firms, and in Keim (1983) that the excess returns of small firms exhibit a very strong January effect. An equal-weighted index gives small firms a greater weight than a value-weighted index.

Lakonishok and Smidt (1988) examine the Dow Jones Industrial Average (DJIA) for the January effect, and do not find January to be a month with above average returns, although they do find some price increases at the end of December that they refer to as the turn-of-the-year anomaly. Since the DJIA is comprised of large stocks, the finding supports the theory that the January effect is limited to small firms. Reinganum (1983) shows that in a given year, losing firms obtain the highest return the following January, supporting the tax-loss hypothesis. On the other hand, DeBondt and Thaler (1985) and Chan (1986) demonstrate that losers continue to experience the January effect up to four years after being identified as losers, which does not support the tax-loss hypothesis.[3]

French (1980) was not the first to demonstrate that Monday's stock return is less than the return on the other days of the week, but he popularized the name and the topic of the weekend effect. Negative returns for Mondays were also observed by Fields (1931), Cross ((1973), and Gibbons and Hess (1981). Because returns after holidays are not unusually low, it is thought that the weekend effect is not caused purely by the market being closed prior to Monday. Subsequent research by Rogalski (1984), using both opening and closing prices, demonstrates that the negative returns accrue over the weekend rather than during the Monday trading day. Possible explanations include the release of negative information over the weekend.

Ariel (1987) examines the daily pattern of returns within each trading month, and concludes that stocks earn significantly greater return in the first half of the month compared to the second half. Ariel includes the last trading day of the month as part of the first half, after noticing that this day is particularly high in returns. Rejecting this classification as data snooping, Lakonishok and Smidt (1989) use the first fifteen calendar days as the first half of the month, and do not find it to be significantly greater than the second half for their sample of the DJIA over a 90 year period.

On the other hand, Lakonishok and Smidt (1988) find a significantly greater rate of return for the four day period from the last trading day of the month through the third trading day of the following month. They call this the turn-of-the-month (TOTM) effect. Ogden (1990) presents and tests a hypothesis that the TOTM effect is caused by the standardization of cash receipts at the end of the month and the re-investment of such payments in the stock market.[4]

The use of new data sets from global markets is beneficial in more ways than one. As mentioned in the introduction, the avoidance of using the same data as everyone else avoids the criticism of selection bias due to data-snooping. But different trading and business cultures, which vary from country to country, can help define the underlying cause of the anomalies. For example, one leading explanation for the January effect is tax-loss hypothesis. But Gultekin and Gultekin (1983) find a January effect in countries that have no capital gains tax.

The TOTM effect has been previously examined in both developed and emerging global markets. Cadsby and Ratner (1992) find a TOTM effect in six out of ten developed countries. Agrawal and Tandon (1994) find a TOTM effect in eleven out of eighteen countries during the seventies and seven out of eighteen countries during the eighties, implying that the effect is diminishing over time. Compton (2002) also finds the effect diminishing in Canada and in the U.S., but strengthening in the Pacific Rim Countries during the nineties.

In this paper we consider the TOTM effect in India using returns on the S&P CNX Nifty, an

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[3] See Keim (1986), Thaler (1987a), and Haugen and Lakonishok (1988) for comprehensive reviews of the early research on the January effect.

[4] Hansel and Ziemba (1996) and Kunkel and Compton (1998) demonstrate the profit potential for exploiting the TOTM by using a switching strategy.

index of stocks trading on the National Stock Exchange. Chotigeat and Pandey (2005) find some evidence of monthly calendar effects using the Bombay Stock Exchange's Sensitivity Index.[5] Nath and Dalvi (2004) examine S&P CNX Nifty index, but focus only for the day-of-the-week effect. We believe we are the first paper to consider the TOTM effect for Indian stock returns.

III. The S&P CNX Nifty

The National Stock Exchange of India (NSE) is one of the largest stock exchanges in the world, based on transactions, and is the largest stock exchange in India.[6] Incorporated in the early nineties, it presently offers trading on debt, equities, equity derivatives, and exchange traded funds. The NSE pioneered the use of an electronic limit order book and internet trading in India.

NSE also set up an index services firm known as India Services & Products Limited (IISL). IISL provides a variety of indices and index related services, and is jointly owned by NSE and CRISIL, India's leading credit rating company. IISL has a consulting and licensing agreement with Standard and Poor (S&P), which allows for the co-branding of IISL products.

One IISL product is the S&P CNX Nifty (Nifty), an equity index of 50 stocks. This value-weighted index is the leading index for large companies trading on the NSE, and is well diversified, representing 25 sectors of the economy. As of early 2006, the Nifty index represented 50% of the traded value and 57% of the market capitalization of the stocks on the NSE. The base value of the Nifty index was defined as 1,000 as of November 3, 1995, although the index was initiated several years earlier. As of July 6, 2006 it stands at 3,156.

IV. Methodology and Findings

The data for our study are the daily returns of the S&P CNX Nifty index from 1992 through 2004. For each trading day we determine the day's position in the month using the following notation: t = 1 is the first trading day of the month; t = -1 is the last trading day of the prior month. We also determine the day-of-the-week of the trading day and ascertain if the trading day falls in January.

After calculating descriptive statistics and the correlations for the sample period, we run the following regression with a dummy variable to compare the returns for TOTM trading day (t = -1 to +3 of each calendar month) and non-TOTM trading days (t = +4 to -2 of each trading month):

$$R_t = a_1 + a_2 TOTM_t + e_t \qquad (1)$$

where:

R_t = rate of return on the portfolio on day t

$TOTM_t$ = 1 if trading day t is at the turn-of-the-month; 0 otherwise

e_t = error term.

[5]The Indian tax year ends in March, and Chotigeat and Pandey (2005) find that returns in March are significantly below returns in January, which they interpret as evidence for a tax-loss-selling hypothesis. But April returns are not significantly positive in an analogous manner to the typical January effect.

[6] This information is based on the NSE web site: http://www.nseindia.com/homepage.htm

Our second regression model adds a dummy variable to capture excess returns if the trading month is January:

$$R_t = a_1 + a_2 TOTM_t + a_3 JAN_t + e_t \quad (2)$$

where:

JAN_t = 1 if trading day is in January; and 0 otherwise.

Our third regression model adds dummy variables for days-of-the-week from Monday through Thursday instead of the January dummy:

$$R_t = a_1 + a_2 TOTM_t + a_3 MON_t + a_4 TUE_t + a_5 WED_t + a_6 THU_t + e_t \quad (3)$$

Where:

MON_t = 1 if the day-of-the-week is Monday; 0 otherwise.

TUE_t = 1 if the day-of-the-week is Tuesday; 0 otherwise.

WED_t = 1 if the day-of-the-week is Wednesday; 0 otherwise.

THU_t = 1 if the day-of-the-week is Thursday; 0 otherwise.

Our fourth regression model adds both the January dummy and the day-of-the-week dummies to the TOTM dummy:

$$R_t = a_1 + a_2 TOTM_t + a_3 JAN_t + a_4 MON_t + a_5 TUE_t + a_6 WED_t + a_7 THU_t + e_t \quad (4)$$

Table 1 shows the descriptive statistics of the daily returns of the Nifty index for the period 1992 - 2004. The first column reports statistics for the entire sample period. The mean (median) daily return was 0.06 (0.06) percent. The second column reports statistics for the TOTM days (-1 to +3 of each calendar month). The mean (median) daily return for the TOTM days is 0.22 (0.28) percent. The third column reports the statistics for the non-TOTM days. The mean (median) daily return is 0.02 (0.03). The final column in Table 1 reports the p-values of the differences between the means, medians, and variances of the TOTM daily returns and the non-TOTM daily returns.

The p-value for the difference between the means are calculated using a two-tailed t-test, and show that the difference between the higher mean returns on the TOTM days and the lower mean returns of the non-TOTM days is significant at the 5% level. The p-value for the difference between the medians is calculated using the non-parametric two-tailed Wilcoxon Scores (Rank Sum) test, and show that the median returns are also significantly higher for the TOTM days at the 1% level. A two-tailed F-statistics test shows that the variances are higher for the TOTM days at the 1% level.

Table 2 shows the correlation measures for the average daily returns on TOTM days and the average daily returns of the non-TOTM days. The correlation between the TOTM returns and the non-TOTM returns is 0.24 with 1% significance.

Table 3 shows the results of all four regression models described above. Newly-West (1987) heteroskedasticity and autocorrelation-adjusted standard errors are used to calculate the reported p-values. Column 1 shows the results for model 1, where the only independent variable is the dummy for TOTM. The coefficient for this dummy is positive and significant at the 5% level. In column 2 (model 2), we report the coefficients and p-values when we add the January dummy to the model. In column 3 (model 3), we add day-of-the-week dummies without the January dummy, and in column 4 (model 4), we add both the January dummy and the day-of-the-week dummies to the TOTM dummy. In all of the regression models above, the TOTM dummy remains significantly positive at the 5% level.

The coefficients for the Monday dummies are negative and significant at the 10% level. The Tuesday dummies are also negative and significant at the 5% level. Both Jaffe and

Table 1: Descriptive Statistics

This table reports the descriptive statistics of daily returns of S&P CNX Nifty index for the period 1992-2004. The first column reports statistics for the entire sample period, the second column reports the descriptive statistics for the turn-of-the-month (TOTM) period (t = -1 to +3 of each calendar month), and the third column reports descriptive statistics for the non-TOTM period (t = +4 to –2 of each calendar month). The final column reports the significance level of differences in mean, median, and variance. The p-value measuring significance of difference in means is calculated using a two-tailed t-test. The p-value measuring significance of difference in medians is calculated using a non-parametric test, the two-tailed Wilcoxon Scores (Rank Sum) test. The p-value measuring significance of difference in variances is calculated using a two-tailed F-statistics.

	R_t, FULL SAMPLE	R_t, TOTM	R_t, Non-TOTM	Difference (p-value)
N	3126	624	2502	
Mean (%)	0.06	0.22	0.02	0.0156
Median (%)	0.06	0.28	0.03	0.0007
Variance	3.08	3.66	2.93	0.0003

Table 2: Correlations

This table reports the correlation measure of average daily returns of S&P CNX Nifty index for the period 1992-2004 during the turn-of-the-month (TOTM) period (t = -1 to +3) and during the non-TOTM period (t = +4 to –2 of each calendar month).

	R_t, TOTM	R_t, Non-TOTM
R_t, TOTM	1.0000	0.2382
p-value		0.0028
R_t, Non-TOTM	0.2382	1.0000
p-value	0.0028	

Westerfield (1985a, 1985b) and Aggarwal and Rivoli (1989) have found a similar Tuesday effect, in addition to the Monday (weekend) effect, in Pacific Rim countries, and have attributed the negative Tuesday returns to the co-movement of global markets. The Far Eastern markets, as well as India, are closed while the U.S. is experiencing the weekend effect of negative Monday returns, which then causes a negative Tuesday trading day in the Far East.

Table 3 also shows that the coefficient for Wednesday is positive at the 1% level. Nath and Dalvi (2004) also find a positive Wednesday effect for their overall sample period. They attribute this to a Wednesday through Tuesday trading cycle, with heavy roll over of previous positions.

The coefficient for the January dummy is not significant in any of the models where it is used. Since the Indian tax year ends in March, we would not expect a January effect based on the tax-loss hypothesis, although it could have been present for other reasons. Chotigeat and Pandey (2005) find a negative January effect for the Bombay Stock Exchange Sensitivity Index.

Table 3: Turn-of-the-Month Effect

This table reports the result of regressions used to calculate the turn-of-the-month (TOTM) average daily excess returns for S&P CNX Nifty index for the period 1992-2004. The dependent variable (Rt) is daily return of S&P CNX Nifty index. TOTM is a dummy variable equal to 1 during the -1 to +3 days of each calendar month. JAN is a dummy variable equal to 1 if the calendar month is January. MON is a dummy variable equal to 1 if day-of-the-week is Monday. TUE is a dummy variable equal to 1 if day-of-the-week is Tuesday. WED is a dummy variable equal to 1 if day-of-the-week is Wednesday. THU is a dummy variable equal to 1 if day-of-the-week is Thursday. Newey-West (1987) heteroskedasticity and autocorrelation adjusted standard errors are used to calculate p-values as reported below the coefficients. *, **, and *** indicate significance at 10%, 5%, and 1% levels.

	Model 1	Model 2	Model 3	Model 4	
Intercept		0.02	0.01	0.05	0.04
	0.6445	0.8280	0.4343	0.5174	
TOTM	**0.20****	**0.20****	**0.20****	**0.20****	
	0.0250	**0.0247**	**0.0284**	**0.0280**	
JAN		0.10		0.10	
		0.4042		0.4077	
MON			-0.17*	-0.17*	
			0.0694	0.069	
TUE			-0.20**	-0.20**	
			0.0354	0.0356	
WED			0.26***	0.26***	
			0.0087	0.0088	
THU			-0.06	-0.06	
			0.4922	0.4900	
N	3126	3126	3126	3126	
Adjusted R^2	0.0018	0.0018	0.0091	0.0091	

V. Conclusion

This paper examines the daily stock returns of the S&P CNX Nifty index, which represents 50 stocks trading on the National Stock Exchange (NSE) of India for the turn-of-the-month (TOTM) effect. To our knowledge, this is the first paper to examine this calendar anomaly for any exchange in India. The NSE is the largest exchange in India and the Nifty index is a value-weighted index of stocks that is a good representative of the NSE stocks.

The TOTM has been studied extensively in the U.S. markets, using essentially the same data base. The advantage of using new data, such as the Nifty index, is that this provides a way to avoid selection bias through data-snooping. In addition, one can discover characteristics of a new market that may not have been previously observed.

We use data for the period 1992-2004, and find a significant difference for the trading days around the TOTM and the non-TOTM. In particular, the returns for TOTM days are greater. In regression models, a dummy variable representing the TOTM is significantly positive, even after controlling for the January effect and for the days-of-the-week effect.

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Mutual Fund Industry: Case Study of Russian Equity Mutual Funds

Natalya V. Delcoure

Assistant Professor in Finance
Department of General Business and Finance
College of Business Administration
Sam Houston State University
Box 2056
Huntsville, TX 77341-2056
Voice (936) 294-1283
Fax (936) 294-3074
e-mail: nvd001@shsu.edu

Dan W. French

Professor of Finance
University of Missouri – Columbia
Columbia, MO 65211
(573) 882-4300
e-mail: frenchdan@missouri.edu

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Abstract:

This paper presents an examination of the mutual fund industry in the Russian Federation. The analysis evaluates the risk-adjusted performance and its relation with the key operating characteristics of Russian open-end equity mutual funds. Risk-adjusted return measures indicate that fund performance is superior to the market. The evidence also suggests that Russian investors should consider fund size and management fees before making investment decisions.

I. Introduction

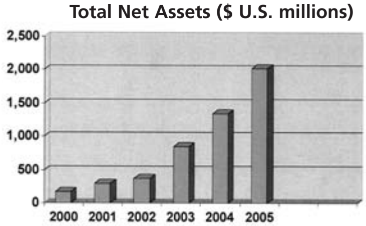
The mutual fund industry in Russia is young (the first mutual fund was launched in 1996) and tiny by U.S. standards. At year-end 2005, the total assets under management barely reached $2.03 billion (Figure 1) compared to about $9.5 trillion in the U.S. This figure is not surprising because until recently the preferred and least risky method of managing savings among Russians was to exchange roubles into dollars or other convertible ("hard") currency and place them under a mattress. After all, in a country where the government destroyed people's savings on more than one occasion through currency devaluations, (winter 1991 and summer 1998), there is little trust in the local currency. But over the past five years, as the currency stabilized and the economy and financial markets grew, Russian investors began looking for more efficient and profitable ways of saving money (Figure 2).

Also, recent changes in the regulatory environment (pension reform of 2002 and 2004, Law on Investing Pension Reserves of 2003, amendments to the Law "On the Securities Market" of 2002) further boosted mutual fund industry growth. By the end of 2005, the number of registered mutual funds in the Russian Federation had grown to 399 (Figure 3). Mutual fund industry growth rate in total assets averaged 32 percent per year between 2000 and 2005. Additionally, in early 2003, the Moscow Interbank Currency Exchange (MICEX) and Russian Trading System (RTS) launched the secondary trading of mutual fund shares making them more appealing to individual and institutional investors, increasing their liquidity, and broadening the geographical range of the client base for the mutual fund industry.

Until recently, most of mutual fund industry expansion has been driven by institutional investors, mainly corporate pension funds, and a few affluent individual investors (Figure 4). More recently, Russia's growing middle class has provided an important source of capital for asset managers.

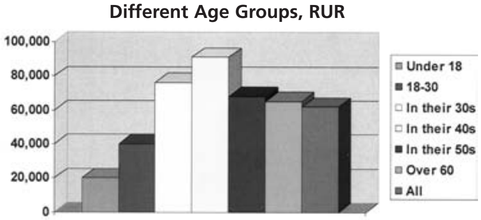
The industry is growing rapidly, and the pension reform and amendments passed by the Russian Parliament – Duma – in 2002 and 2004 are likely to propel the industry forward in a significant way. Even with the influx of retirement money, the industry still faces the substantial task of educating Russian investors by publicizing the merits of mutual fund investing and showing how a private company (mutual fund) can manage retirement savings to yield a greater return than the State Pension Fund. In Russia, the current average pension is only $63 a month. Most Russians do not even understand the basics of mutual fund investing, and asset management is still an entirely new and rather fragmented industry.

Figure1: Russian Mutual Fund Industry Total Net Assets ($ U.S. millions)



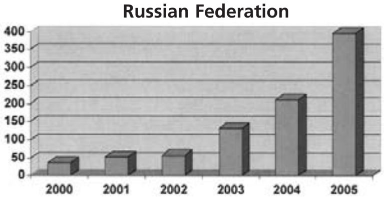
Source: Investment Company Institute, 2005: http://www.ici.org

Figure 2: Average Mutual Fund Purchase across Different Age Groups, RUR



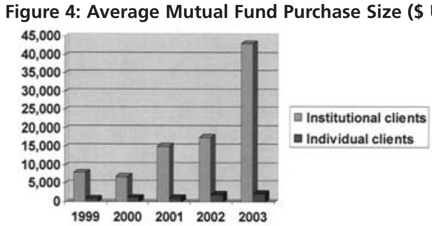
Source: PIOGLOBAL Company data, 2003: www.pioglobal.com

Figure 3: Number of Mutual Funds in Russian Federation



Source: Investment Company Institute, 2005: http://www.ici.org

Figure 4: Average Mutual Fund Purchase Size ($ US)



Source: PIOGLOBAL Company data, 2003: www.pioglobal.com

The purpose of this paper is to investigate the mutual fund industry in the Russian Federation. Russia possesses one of the largest emerging equity markets in the world, with a total market capitalization of approximately $285 billion at year-end 2005. To preview the results, we note that open-end equity mutual fund risk-adjusted return is superior to broad equity market returns in the period between 1999 and 2005. Also, empirical evidence suggests that Russian investors should follow conventional wisdom and consider fund size and management fees before making investment decisions.

II. Financial markets in the Russian Federation

In the early 1990's, Russia's financial market situation was bleak. The country's budget deficit was high, and the government had little ability to collect taxes. The 1998 foreign debt crisis and subsequent currency devaluation made it virtually impossible for the Russian government to acquire new funds. The Russian government bond market was closed down and trade on the MICEX was halted more than once. Lack of confidence in the Russian financial system led to runs on banks, prompting the government to guarantee household deposits in the State savings bank, though not in private banks. The economy slid into depression, the Russian stock market had to be shut down on numerous occasions, and foreign investors withdrew their investments from the Russian Federation in a "flight for quality." The 1998 financial collapse was due to deficit spending that was unsustainable without causing hyperinflation and currency devaluation.

In the immediate aftermath of the financial crisis, the government undertook efforts to address persistent weaknesses in the investment climate and tackle the fiscal and monetary challenges of the Russian economy. Principal among these efforts was the adoption and quick implementation of a simplified tax policy (including a flat tax rate of 15%) which increased government revenue, and the amendment to the law "On the Securities Market." This law provided definitions of corporate bonds, mutual funds, derivative securities (options, futures, and forwards) affording a sounder legal basis and encouraging increased transparency in the financial markets.

Since 2003, companies offering public shares are required to disclose financial and accounting information during the placement process as well as in their subsequent quarterly and annual reports. In addition, the new Law clearly defines "price manipulation" as well as responsibilities of investment bankers assisting companies with their securities offering. However, the amendment does not cover "insider trading," the subject of another bill stalled in the Duma in late 2006 as a result of strong lobbying against its passage. In 2004, the Duma passed the law "On Mortgage Securities" as part of a housing reform package in an effort to facilitate the accessibility of long term funds used for mortgage lending. The law provides for the development of a mortgage market, the issuance of mortgage-backed securities, and protections for individual investors.

Meanwhile, the executive branch of the Russian Government showed fiscal responsibility, applying 100 percent of the budget surplus to the country's debt repayment. Rising commodity prices helped to reduce the budget deficit even further, boosting Russian gross domestic product, foreign reserves and citizens' disposable income. The Russian banking industry and financial markets began to acquire many of the technical and functional characteristics of mature markets.

Despite their recent growth, both the private banking sector and financial markets remain small relative to the country's GDP. For example, in 2003 bank credits financed only about five percent of fixed investments in Russian economy vs. 94 percent in the Czech Republic. Fifteen years after the emergence of the first commercial bank in Russia, the banking sector still remains small, fragmented, and dominated by a few mostly state-owned institutions even though formal barriers to foreign banks' entry have been relaxed significantly. Ownership structure in banking continues to remain obscure, and there is little trust among banks or between banks and clients unless they have ownership or other ties. The inter-bank market is underdeveloped, and there is little of the interaction among banks normally found in a well functioning network of financial intermediaries. There is little pooling, trading and risk sharing.

Russian non-bank financial institutions, while expanding rapidly over the last five years, remain at an early stage of their institutional development. At year-end 2005, twelve stock exchanges operated in Russia. The RTS and the equity trading floor of the MICEX are the dominant exchanges. Although the number of companies traded on the RTS is greater compared to the MICEX (700 issuers vs. 170), average trading volume on the MICEX is much higher than on RTS ($205.54 billion vs. $225.6 billion in 2005). Several Russian regional centers have their own stock exchanges, but trading volume outside Moscow tends to be low.

While total stock market capitalization in 2005 is relatively high for an emerging market (around $285 billion ~ 50 percent of country GDP), indicators of liquidity and market depth are still poor. Although the market boomed for most of 2001 - 2005, there have been only a few initial public offerings. As a result, the shares of only fifteen companies (mainly oil, gas, steel, and telecommunication) account for 90 percent of market capitalization and turnover on the RTS and the MICEX. Many large Russian corporations (e.g., Gazprom, LUKoil, Tatneft) listed on the RTS and the MICEX continue to issue American Depository Receipts (ADRs) on the New York Stock Exchange (NYSE).

Since 2002, the corporate bond market has become an increasingly important alternative to bank financing. The Russian corporate bond market is segmented into rouble- and hard-currency debt. In 2004, the market increased significantly in size – 71 companies and 16 banks issued bonds compared to a total of 60 issuers in 2003. In addition, in 2004 15 corporations and 10 financial intermediaries issued Eurobonds in the amount of $12 billion. The reason for such rapid growth of the corporate bond market is the absence of an effective banking mechanism of lending funds to enterprises, especially to smaller companies.

Somewhat larger than the corporate bond market is the market of more popular unsecured promissory notes with specified maturities ranging from six months to two years. Called veksels, these notes are issued by companies, banks, and the government. Veksels are sold at a discount in the over-the-counter market with the major banks acting as market makers, and they may or may not have a coupon payment. Though veksels are popular as a liquidity management instrument and are less expensive to issue, they are riskier investments compared to corporate bonds because their underwriting requirements are less demanding and their legal status is less well defined. Nevertheless, at the end of 2003, there were about $10 billion of veksels outstanding.

While the corporate bond market is rapidly developing, it remains quite narrow. The market has been unable to provide satisfactory liquidity for small issuers, even if they are "blue-chip" companies. Also, the underwriting cost and stamp duty (a 0.8 percent tax) add to expenses. In 2005, the volume of placements amounted to only one percent of GDP, and the role of the Russian bond market as an alternative source of finance remains modest. Another barrier to bond market growth is a provision of Federal law "On Joint Stock Companies" which requires the volume of bond financing not to exceed a company's charter capital. Thus, the domestic financial market remains a limited source of investment capital for corporations.

The government and municipal debt market also is underdeveloped. At year-end 2004, the total volume of rouble-denominated government debt outstanding was about five percent of GDP. Only 40 percent of this debt was tradable, and a very large share of that was held by state-owned banks.

III. The mutual fund industry in Russian Federation

Mutual funds have existed and effectively operated in developed markets for more than 80 years. The first mutual fund in the Russian Federation was formed in 1996, and as the stock market boomed the number of funds proliferated until the financial crisis in 1998 killed all but one mutual fund. Since the early 2000;s, the number of funds has rebounded and is growing as a result of Russia's strong economic development, rising incomes, and recently passed State Pension Fund reform. At year-end 2005, there were 120 asset management companies in Russia operating 209 open-end, 127 closed-end, and 63 interval active mutual funds (Figure 5). However, 60 percent of all assets under management is concentrated within the four largest companies (Management-Center, Management-Consulting, UralSib, and Nikoil-Savings).

Barriers to entry in the mutual fund industry have been low compared to developed countries. For example, the minimum equity capital requirement is $80,000 for an open-end mutual fund. At the same time, regulators impose constraints on eligible asset classes allowing investment only in the most liquid domestic stocks, and no more than 20 percent of the portfolio in foreign securities. These requirements limit security selection and diversification choices. The relatively short list of permitted securities is considered to be exhaustive: no one can create new financial instruments without legislation. The industry is in desperate need to make existing securities law more transparent and open to financial innovations. In particular, there is a need to provide a stronger legal basis for such activities as financial engineering and asset securitization.

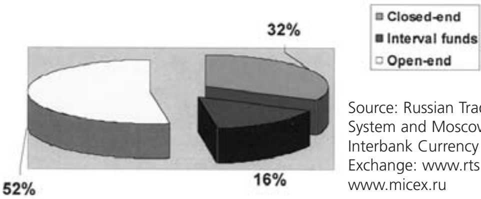
Furthermore, there is no proper legislation governing derivative securities. Small markets in derivatives instruments exist, but their legal position has been uncertain. A Russian court ruled that derivatives contracts fell under the Civil Code provisions pertaining to wagers and so were not subject to court enforcement.

Currently, the mutual fund industry in Russia faces significant challenges. The first challenge is finding companies that meet investment and liquidity standards. Koke (1999) determines that a reduction in general, microeconomic, and financial risk factors increases investments of Western mutual funds in Central and Eastern European (CEE) equity and bond markets. He shows that Western institutional investors check a whole range of factors (e.g., the stability of legal, economic, and financial systems; currency stability; strictness of bankruptcy law; liquidity of financial markets; presence of strategic investors; level of taxes, and others) before buying assets in CEE countries.

On the microeconomic side, investors look for highly qualified management teams. Frydman et al. (1997) suggests that competent decisions regarding capital and labor allocation are particularly important in the former centrally planned economies. Also, Blommestein (1998) finds that institutional investors have a strong preference for liquid assets. In Russia, these preferences may be of particular relevance because some financial instruments are very thinly traded and are hindered by the underdevelopment of clear and fair legislation protecting financial markets.

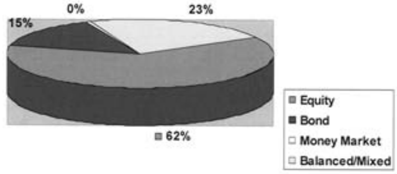
In western countries, the activity of mutual funds is generally subject to strict regulation. In the Russian Federation, the Ministry of Finance, the Bank of Russia, and subsequently the Federal Commission for the Securities Market (FSM) are the main regulators. In January 1996, the participants of the government bond market founded a self-regulatory organization called the "National Securities Market Association" (NSMA). In December 2003, the NSMA became an associate member of the International Securities Market Association (ISMA). At present, NSMA is the main professional association in The Russian Federation and has the full set of government licenses to carry out the supervision, analysis of business reporting information and inspection functions over banks and investment companies regarding their brokerage services, depositary and trust management activity. At the same time, the FSM established state control for mutual

Figure 5: Russian Mutual Funds, 2005



Source: Russian Trading System and Moscow Interbank Currency Exchange: www.rts.ru, www.micex.ru

Figure 6: Total Net Assets in $U.S. by Type of Fund, 2005



Source: Investment Company Institute, 2005: http://www.ici.org

fund activities, separated the control of assets of a fund from their administration, organized a multilateral cross check of mutual funds activities, and identified a set of information disclosure necessary to increase investor confidence.

Figure 6 classifies Russian mutual funds by type of investments. It appears that Russian investors prefer equity mutual funds to bond, balanced/mixed, and money market mutual funds. In fact, money market and index mutual funds remain a novelty on the Russian market. Since 2004, Russian investors have had a choice among only three money market and eight index mutual funds.

The minimum investment for a retail investor in a mutual fund is fairly low – 1,000 roubles (~ $30) for an open-end and 5,000 roubles (~$150) for a closed-end mutual fund. One of the interesting features of mutual fund investing is that Russian investors generally have a short investment horizon (typically less than six months). Also, they switch from one fund to another more frequently compared to U.S. investors, treating mutual funds more like individual stocks. Furthermore, according to PIOGLOBAL research, 47 percent of Russians make investment decisions emotionally rather than rationally. It appears that Russian investors' short-term investment horizon is a matter of culture and inadequate knowledge about the mutual fund industry. Grigoriev, et al. (2003) report that in 2001, when the Russian market grew most rapidly, only 3% of the growth of mutual fund industry could be attributed to the inflow of money from new investors. Presently, asset management companies in Russia spend millions of dollars on marketing and advertising campaigns aimed at both retail and institutional clients.

Another obstacle to industry growth is the lack of a distribution network. Since very few "stand-alone" asset management companies have a retail branch network in the geographically spread Russian Federation compared to their hybrid bank/funds competitors, selling mutual funds is a challenge. It appears that the asset management business in Russia is still more about education rather than market share acquisition.

IV. Data and Methodology

To investigate the performance of Russian mutual funds, we collected net asset values (NAVs) for open-end mutual funds from January 1999 through December 2005 from www.kommersant.com the website of Kommersant – the first independent newspaper in the Russian Federation. The sample includes 120 open-end mutual funds which are listed on the MICEX or the RTS and have closing NAVs available. Open-end mutual funds account for 52 percent of the mutual fund population (Figure 5). Shares of these funds are bought and sold on a daily basis. Another reason we restrict this analysis to open-end mutual funds is that even though closed-end mutual funds comprise 32 percent of the active mutual fund industry, they allocate cash to "special purpose" projects (e.g., commercial real estate). Finally, interval mutual funds are open for transactions only several times a year. These financial intermediaries usually invest in equity of fast growing companies although most of them suffer from low market liquidity. Also, interval funds re-compute their NAVs once a month at best, and more than 62 percent of total net assets of existing interval funds belong to only four Nikoil Asset Management interval funds.

We adopt the six-year data period from 1999 to 2005 which starts after the latest Russian currency and financial crisis through the time for which complete data are available. The paper uses the $U.S. exchange rate and RTS market index, which is the capital weighted index, calculated based on 50 of the most liquid and highly capitalized Russian companies.

The sample is free of survivorship bias. Survivorship bias is a well-documented problem in mutual fund research since it may overstate average performance measures resulting from the exit of poorly performing funds through liquidation or merger. This study reduces the survivor problem by collecting time series data each year from all of the funds in existence for a given year. The final sample contains data of all equity funds listed for a given year and includes surviving funds as well as funds that cease to exist.

Another challenging problem of the Russian data is the absence of risk-free assets. After the currency crisis in August 1998, it is difficult to view the existing government debt instruments as risk-free investments. We follow the Barinov (2003) approach and use the overnight rate for loans offered by the Central Bank of Russia as a proxy for a risk-free interest rate. Finally, mutual fund operating characteristics (e.g., asset size, expense ratios, management and custodian fees, and load/no-load status) are collected manually from the funds' websites.

The sources of mutual fund management fees differ from funds in the U.S. market. In Russia, there are two main categories of costs borne each year by mutual fund investors. First is the management fee paid every year to the management company as a percentage of the fund's total assets. Such fees vary from 0.99% to 15.1%. Second are other fees that include custodial fees, trading costs, the stamp duty, and brokerage and auditor fees. These fees range from 0.9% to 5.64% of fund total assets.

IV.1. Performance Measurement Models

This study investigates the performance of the mutual fund industry in the Russian Federation using the sample of open-end equity mutual funds. Starting with Sharpe (1966) and Jensen (1968), academic studies have been dedicated to analysis of the investment performance of mutual funds. Some of these studies examine mutual fund performance relative to a return benchmark. Others investigate differences in mutual fund performance and attempt to explain their causes and consequences.

First, we examine the extent to which actively managed open-end equity mutual funds in Russia earn excess risk-adjusted returns compared to the Russian equity market returns (as measured by the RTS index). Fund performance is measured using the standard Jensen (1968) equation:

$$R_t - R_{ft} = \alpha + \beta * (R_{mt} - R_{ft}) + \varepsilon \qquad (1)$$

where

R_t – unadjusted NAV return for the fund in time t;

R_{ft} – Central Bank of Russia overnight lending rate for time t;

R_{mt} – return on the RTS index for time t.

Second, we examine the extent to which mutual fund investment performance is related to key operating characteristics of the fund. The key operating characteristics are the fund size (measured by total assets), expense ratio, load/no-load status, and management and custodian fees.

Previous mutual fund research is inconclusive on the interrelationship between the fund size, cost variables (e.g., loads, operating expenses, asset turnover, and management fees) and fund performance. The effect of scale on mutual fund performance is an important question. Some practitioners point out that large funds have more resources to spend on research, hire more managers to generate additional good ideas, and lower expense ratios. On the other hand, Carlson (1970) and Volkman and Wohar (1995) show that fund size is not related to performance. In contrast, Grinblatt and Titman (1989) and Gorman (1991) find negative relations between gross returns and fund size. Sharpe (1966) finds that funds with lower expenses realize superior performance. However, Ippolito (1989) reports no significant relationship between fund performance, investment fees and turnover. Golec (1996), Hooks (1996), and Dellva and Olson (1998) conclude that fees and load expenses are associated with negative mutual funds' excess returns. At the same time, Droms and Walker (1994) uncover a positive relation between fund returns and expenses and no relation between performance and loads.

More recently, Khorana and Nelling (1997) find no effect of management fees, turnover ratio and load on sector-specific mutual fund returns. Prather et al. (2004), after considering general market conditions and a fund's investment objective, determine that fund performance is negatively related to fund size and expense ratio. Furthermore, Cesari et al. (2002) find that Italian equity fund performance is negatively related to mutual fund investment fees.

Thus the existing literature often reaches conflicting conclusions regarding the relationship between mutual fund performance and its operating characteristics. The present study builds on the previous empirical research to provide the first examination of Russian equity mutual funds' performance.

The following regression model examines the interaction between mutual fund performance and key operating characteristics:

$$R = f(A, E, C, L) \qquad (2)$$

$$R = \hat{a}_0 + \hat{a}_1 A + \hat{a}_2 E + \hat{a}_3 C + \hat{a}_4 L$$

where

R = annualized NAV rate of return. We use the following three measures of annual risk-adjusted rate of return in the regression analysis: the Jensen (1966), the Sharpe (1966), and the Treynor (1965) measures.

$\hat{a}$ = regression estimates;

A = natural log of mutual fund Total Assets;

E = fund management expenses as a percentage of NAV;

C = fund custodian fees as a percentage of NAV;

L = zero for no-load funds with no sales charges and one for load funds.

V. Discussion of results

Table 1 presents means and standard deviations for the market index, open-end mutual fund returns and their key operating characteristics in Russian Federation along with the tax rate on capital gains and dividends.

Table 2 summarizes the mutual fund risk-adjusted performance measures and compares them to the market portfolio. The estimated alpha coefficient is positive and statistically significant. It suggests that open-end mutual funds outperformed the market between 1999 and 2005. The beta coefficient is less than one and statistically significant. Overall, open-end mutual funds provide diversification benefits to Russian investors without sacrificing rate of return. The explanatory power (R2) of the Jensen (1968) model is .61 suggesting that the market variable explains more than one-half of the variation of the equity mutual fund performance.

The Sharpe (1966) and Treynor (1965) risk-adjusted performance measures are higher for the sample of open-end mutual funds compared to the market portfolio. Both risk-adjusted return measures consistently show that actively-managed mutual funds outperformed the market during the sample period.

Table 3 provides the results for cross-sectional estimation of the relationships between mutual funds' risk-adjusted rate of return and their key operating characteristics. None of the intercept coefficients is statistically significant. The performance of Russian open-end mutual funds depends on the size of assets under management and management expense ratio. The regression results suggest that the asset growth is positively related to the performance of Russian open-end equity mutual funds.

Our empirical findings do not support the expectation that operating and management expenses of Russian mutual funds support research, marketing, and managerial expertise. The result for management expenses is consistent with Ippolito's (1989), Droms and Walker's (1994), and Malhotra and McLeod's (2000) results.

The load/no-load estimated coefficient shows no difference in the performance of load versus no-load mutual funds. These results are consistent with the previous findings by Droms and Walker (1994) and Elton et al. (1991). Finally, custodial fees do not exhibit a significant effect on the open-end equity mutual fund performance in Russia.

This finding is somewhat puzzling. In an efficient market, these fees should be a direct "drag" on mutual fund returns. We speculate that the observed relationship may be explained by Russian market inefficiencies.

Table 1. Summary statistics on fund characteristics

Variable	Mean	Median	S.D.
General Information			
Market return, %	13.25%	9.77%	10.52%
Mutual fund return, %	16.82%	12.84%	17.84%
Total fund assets ($M)	$569.17	$316.09	$176.7
Cost Variables, %			
Management expenses	9.16	7	7.12
Custodian fee	3.41	2.49	0.52
Tax rate on capital gains and dividends	15% flat tax rate		

Table 2. Risk-adjusted mutual fund performance measures

	Jensen (1968) measure	Sharpe (1966) measure		Treynor (1965) measure	
		Open-end mutual funds	Market portfolio	Open-end mutual funds	Market portfolio
		0.79	0.47	25.48	8.56
α	0.38* (0.085)				
β	0.56*** (0.002)				
R2	0.61				

Note: p-value reported in parenthesis. * Significant at 0.1 level. ** Significant at 0.05 level. *** Significant at 0.01 level.

Table 3. Mutual fund risk-adjusted performance and key operating characteristics

	Constant	Ln(TA)	Management expenses	Custodian fees	Load dummy	R2
Jensen measure	-0.82 (0.91)	0.10** (0.01)	0.01 (0.12)	0.70 (0.71)	0.60 (0.52)	0.53
Sharpe measure	-0.35 (0.20)	0.21** (0.01)	-0.10* (0.06)	0.86 (0.50)	0.67 (0.93)	0.46
Treynor measure	-0.28 (0.64)	0.09* (0.07)	-0.001* (0.02)	0.90 (0.62)	0.61 (0.84)	0.44

Note: p-value reported in parenthesis. * Significant at 0.1 level. ** Significant at 0.05 level. *** Significant at 0.01 level.

VI. Conclusion

The purpose of this study is to provide the first comprehensive study of the performance of Russian equity mutual funds. The empirical results are free of survivorship bias. The analysis evaluates the risk-adjusted performance of Russian open-end equity mutual funds and the relations with their key operating characteristics. With risk-adjusted returns, the funds' performance is superior to the market. Also, the empirical evidence suggests that Russian investors should consider funds' size and management fees before making investment decisions.

The pension reform in Russian Federation opens up a new chapter in the Russian securities market. We speculate that pension capital infused by individual investors through asset management companies will play a major role in the future growth of the Russian securities market and mutual fund industry. Given the enormous literature analyzing mutual fund performance and the impact of funds' operating characteristics on the persistence of performance, further research should investigate whether Russian equity funds' risk-adjusted returns are stable over time as the data becomes more available.

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The Dynamics of Volatility Spillovers in Chinese Futures Markets

Leo H. Chan*

Assistant Professor of Finance
Delaware State University

Donald Lien

Professor of Economics
University of Texas - San Antonio

Wai-Kin Leung

Associate Professor of Finance
Hong Kong Chinese University

Abstract:

This paper investigates the relationships among volatility series in three major futures markets (mungbeans, soybeans and copper) in China, with an attempt to identify possible cross-market spillovers. Using VAR approaches augmented by open interest and trading volume data, we find a bi-directional relationship prevails between soybeans and mungbeans futures markets and the copper market dominates both mungbeans and soybeans markets in terms of volatility transmissions. Results from Geweke measures suggest bi-directional relationships among the three volatility series with instantaneous feedback measure being the strong dominant factor. In a more comprehensive VAR framework incorporating lagged volume and open interest data from each market, we detect no volatility spillovers across the three markets. Instead, cross-market effects are likely to prevail between the trading volume in one market and volatility in another.

JEL Code: G14
Key Words: Chinese Futures Markets, Volatility Spillover

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Corresponding author: Leo Chan, Department of Accounting and Finance, School of Management, Delaware State University, 1200 N. DuPont Highway, Dover, DE 19901. Tel: (302) 857-6919, Fax: (302) 857-6924, e-mail: lchan@desu.edu

I. Introduction

The functions and effectiveness of derivatives markets in a developed capital market have been well established. The hypothesis stated in Ross (1976) has been tested and confirmed by many empirical studies. Due to, in large part, lack of data, studies of derivatives markets in emerging economies have been sparse. The challenge faced by researchers in dealing with derivatives markets in emerging economies goes beyond lack of data when the underlying spot markets are relatively young (or missing all together). Such is the case for China's futures markets. The lack of organized spot markets, and thus lack of spot market data, means traditional test of price discovery function and hedging effectiveness cannot be performed. Chan, Chan and Leung (2005) investigate the effect of Zhengzhou Commodity Exchange's (ZCE) decision to switch focus from the mungbeans futures contract to the wheat futures contract and found strong impacts on the volatility of both contracts. Chan, Fung and Leung (2004) study the impact of regulatory reform on futures markets volatility and found significant reductions in volatility in all existing futures markets. The results in both studies point to a system of futures markets that is influenced mainly by regulatory and institutional forces. Chan et al. (2005) suggest that futures markets in China could be dominated by speculators rather than by hedgers. This paper extends the finding in Chan et al. (2005) by investigating the transmission of information across three major commodity futures markets in China. More specifically, we attempt to investigate whether there are significant information transmissions between these futures markets and if there are common factors that drive volatility in these markets.

II. Literature Review

China's futures markets are relatively young and the market development and changes have been rapid since the inception. In 1993, the first futures contract for mungbeans was traded on ZCE. By the end of 1995 there were fourteen organized exchanges in China. Lien and Yang (2004) provide a brief history of the growth and development of futures markets in China and a summary of empirical findings in the literature. Williams, Peck, Park and Rozelle (1998) document the development of the first successful futures market in China – the ZCE. The exchange started with great expectations, introduced over 11 contracts (including many commonly used agriculture products such as rice, soybeans, peanuts, and corn) for trading in the first three years. By mid-1995, all but wheat and mungbeans futures are de-listed due to low trading volume.

Chan, Fung and Leung (2004) describe major institutional changes in the development of China's futures markets. To standardize trading and centralize regulatory reforms, the Chinese government combined the fourteen futures exchanges in 1998. As a result, all but three, the Shanghai Futures Exchange (SFE), the Dalian Commodity Exchange (DCE) and the Zhengzhou Commodity Exchange (ZCE), remained. In each of these three exchanges, only one futures contract has consistent good trading volume at any given period.

With advice from Chicago Board of Trade, the ZCE developed a wholesale market for mungbeans prior to the establishment of the futures market. This practice is not uncommon as the existence of an active spot market is beneficial for the functioning of the corresponding futures market. Chan et al. (2005) suggest that the lack of an active spot market for mungbeans might have led to the declining interest in the mungbeans futures trading. In late 1999, the ZCE started to promote wheat futures contracts, starting with the January 2000 contract. The promotion effort led to a significant increase in the trading volume of the wheat futures market, at the cost of a great reduction in the trading volume of the mungbeans futures market. The wheat futures contract replaces the mungbeans futures contract as the most actively traded contract in ZCE in a matter of days following the promotional effort by the exchange.

These studies all investigated the futures markets in China separately. In a developing market where participants are learning the functioning of the system, it is likely that there is cross-market learning as well. Our paper contributes to the existing literature and understanding of the development of financial markets in emerging markets in that we investigate whether there are significant information transmissions between the futures markets in China, and if there are common factors that drive volatility in these markets.

III. Data and Methodology

Due to lack of spot market data, we can only use data from the futures markets to study information transmissions between these markets. We examine three futures contracts, one from each exchange, namely, mungbeans (ZCE), copper (SFE) and Soybeans (DCE). Note that although the ZCE has two active contracts (mungbeans and wheat) during much of the period covered in our data set, Chan et al. (2005) find that the volume data for the wheat contract is not reliable prior to 2000. Thus, we utilized only the mungbeans contract from ZCE in this paper. The period under consideration is between January 5, 1998 and November 29, 1999. After adjusting for difference in trading days, we have 467 observations for each contract.

To study the interaction of volatilities among the three futures markets, we first construct range-based volatility measures based on the daily high, daily low, daily open and daily close prices proposed by Alizadeh, Brandt, and Diebold (2002), and Rogers and Satchell (1991). The bi-directional feedback measures suggested by Geweke (1982) are then computed to analyze the interactions between these markets via three pair-wise comparisons. The vector autoregression (VAR) approach recommended by Chan and Cheung (1993, 1995) is also applied to investigate the relationships among open interest, trading volume, and volatility.

The construction of the range-based volatility is as follows: Consider a trading day period, denoted by t. Let Ht, Lt, Ot and Ct denote, respectively, the high, low, open and close prices of a given futures contract at day t. The volatility measure, Range t , is defined as the difference between the high and low prices (in logarithms):

$$Range_t = \ln(H_t) - \ln(L_t) = \ln(H_t / L_t) \qquad (1)$$

Gallant, Hsu, and Tauchen (1999) and Alizadeh, et al. (2002) both found range to be an information-rich (low variance) proxy for the true volatility.
For range to be a valid volatility measure, one must assume the futures price follows a random walk with zero drift. This assumption is most likely to be falsified in most financial data. To remedy this drawback, Rogers and Satchell (1991) proposed an alternative measure that is drift-independent:

$$VRS_t = [\ln(H_t) - \ln(O_t)][\ln(H_t) - \ln(C_t)] + [\ln(L_t) - \ln(O_t)][\ln(L_t) - \ln(C_t)]. \qquad (2)$$

The resulting volatility measures from (1) and (2) are: mungbeans *(RangeM)*, copper *(RangeC)*, and soybeans *(RangeS)* or *VRS* measures for mungbeans *(VRSM)*, copper *(VRSC),* and soybeans (VRSS). After we obtained the two volatility measures in each of the three futures markets, we fitted each volatility series with a pair-wise VAR model to generate Geweke measures. Details of how the Geweke measures are constructed can be found in Chan, Chan and Cheng (2004).

In addition to interactions between volatility series of the futures markets, we also examine interactions between open interest, trading volume and volatility across the three markets using the comprehensive VAR approach suggested by Chan and Cheung (1993). With three volatilities *(RangeM, RangeC* and *RangeS,* or *VRSM, VRSC* and *VRSS*), three open interests *(OIM, OIC* and *OIS)* and three trading volumes *(VOLM, VOLC* and *VOLS)*, the comprehensive VAR model we utilize in this paper consists of a system of nine equations[1].

1 To conserve space we did not write out the system explicitly. Interested readers can contact the author for details.

IV. Empirical Results

Summary statistics for both volatility measures show that mungbeans futures market has the largest average volatility, followed by soybeans and copper. The same observations apply to higher moments of volatility series. Note that, during our sample period, the importance of the mungbeans futures market had diminished with the improved trading volume of the wheat futures market. On the other hand, the copper futures market has a long history and is widely believed to be a hedger's market.

Tables 1 and 2 present the results from the structural equations (the first required step for construction of Geweke measures) that capture potential cross-market interactions. Table 1 reports the estimation results for pair-wise relationships among these three contracts, augmented by own open interest and trading volume, using the range-based volatility measure. Table 2 reports the results of the pair-wise relationship using the Rogers-Satchell volatility measure. For the copper and mungbeans markets, both measures indicate that the volatility of the copper futures market affects the volatility of the mungbeans futures market whereas the latter has no effect on the former. These results imply a unidirectional effect from the copper market to the mungbeans market. However, the range data suggests a positive impact but the VRS data predicts a negative impact. Using either volatility measure, the volatility in either market increases with increasing own trading volume and decreasing own open interest. That is, trading volume is associated with diverse information, leading to greater price fluctuations, and open interest is related to hedge positions and tends to stabilize the futures price. Our results are consistent with the findings of other empirical studies that utilized open interest as proxy for market depth and trading volume as proxy for speculative activities (see Fung and Patterson, 2001).

For the mungbeans and soybeans markets, because both mungbeans and soybeans are a main staple food in central and northern China, we expect the volatility to spillover across the two markets. Using both volatility measures, we confirm the above conjecture. That is, there are bi-directional information flows between mungbeans and soybeans futures markets. Once again, we find that the volatility in either market increases with increasing own trading volume and decreasing own open interest.

Turning to the results between copper and soybeans markets, we have similar observations on volatility spillovers to that between copper and mungbeans markets. In addition, the effects of own trading volume and own open interest are the same as the previous two cases. There is a unidirectional effect from the copper market to the soybeans market. Summing information in Tables 1 and 2, we conclude that a bi-directional relationship prevails between soybeans and mungbeans futures markets whereas the copper market dominates both mungbeans and soybeans markets in terms of volatility transmissions.

The Geweke measures reported in Table 3 seem to provide a different story. In each pair of futures markets, the instantaneous feedback measure is the strong dominating factor. The largest values occur in the soybean-copper pair (5.9437 for VRS and 10.69697 for range). The smallest values prevail in the soybeans-mungbeans pair (3.7468 for VRS and 9.66389 for range). Similar orderings apply to the total feedback measures. Interestingly, in each pair, we detect bi-directional information flows. The lead-lag relationships between copper and mungbeans (and soybeans) based upon the previous VAR method are now non-existent. The difference is likely to arise from the addition of own trading volume and own open interest in the VAR models. Using the results from both volatility measures, we find the unidirectional feedback measure is smallest from mungbeans to copper, followed by the case from mungbeans to soybeans. Thus, the mungbeans futures market appears to have the lowest degree of integration with the other two markets. Once again, this result may be attributed to the fact that, during the sample period, the ZCE was contemplating to shift its focus from the mungbeans futures market to the wheat futures market.

To investigate if there are interactions between open interest, trading volume and volatility across the three markets, we turn to the comprehensive VAR results in Table 4 and 5. We are mainly interested in the cross-market interactions. For the volatility series, only the trading volume of the mungbeans market has a statistically significant (positive) effect on the volatility of the copper market. The volatility of the copper market produces a negative impact on the trading volume of the mungbeans market and a positive impact on the trading volume of the soybeans market. The above results reinforce the concept that open interest is associated with the hedge position whereas volatility and trading volume are both tied to speculative activities generated by diverse information. The relationships among various statistics in own market show that open interest leads the trading volume in each market. In addition, market volatility leads the trading volume in copper and soybeans markets whereas a two-way feedback relationship between volatility and trading volume in the mungbeans market. There is no intertemporal relationship between open interest and volatility in each market.

Stronger cross-market interactions are observed in Rogers-Satchell measure reported in Table 5. The trading volumes of soybeans and mungbeans markets both have impacts on the volatility of the copper market. The trading volume of the soybeans market is positively affected by the trading volumes in the two other markets, whereas the trading volume in the mungbeans market is affected by the volatility and trading volume in the copper market. Once again, there is no cross-market impact on open interest series. Finally, the volatility of the copper market is affected by the trading volumes in both mungbeans and soybeans markets. In addition, the volatility of the soybeans market is affected by the volatility of the mungbeans market. From the trading volumes side, there is interaction from mungbeans to soybeans, and from copper to mungbeans and soybeans futures. The results indicate stronger speculative activities across the three markets.

In terms of the relationships among various statistics in own market, the results in Table 5 are similar to that from Table 4. It is worth noting that there is no cross-market interaction among any of the open interest statistics. Since open interest is a proxy for market depth, the lack of cross-market interaction from this statistic suggests that information from true users have no impact on the activities of other markets. This finding is similar to the conjecture suggested by Chan et al. (2005).

V. Conclusion

This paper investigates the relationships among volatility series in three major futures markets in China, with an attempt to identify possible cross-market spillovers. Using VAR approaches augmented by open interest and trading volume data, we find a bi-directional relationship prevails between soybeans and mungbeans futures markets and the copper market dominates both mungbeans and soybean markets in terms of volatility transmissions. Geweke measures suggest bi-directional relationships among the three volatility series with instantaneous feedback measure being the strong dominant factor. In a more comprehensive VAR framework incorporating lagged volume and open interest data from each market, we detect no volatility spillovers across the three markets. Instead, cross-market effects are likely to prevail between the trading volume in one market and volatility in another. Thus, we conclude that there is no direct volatility spillover among the three Chinese futures markets. Indirect spillovers through trading volume and open interest do prevail. When the two mediating variables are not taken into account, the resulting Geweke measures attribute indirect effects to direct effects. Nonetheless, while these lead-lag effects are statistically significant, they are very small compared to the instantaneous feedback effects.

Table 1: Structural Equations for VRS Measure

VRS	Soybean	Mungbean	VRS	Copper	Soybean	VRS	Copper	Mungbean
Constant	0.0437**	0.1186**	*Constant*	0.0147**	0.0578**	*Constant*	0.0155**	0.1792**
$VRSS_{t-1}$	0.1868**	-0.1055	$VRSC_{t-1}$	0.1730**	-0.2550	$VRSC_{t-1}$	0.1586**	-0.6151
$VRSS_{t-2}$	0.0278	-0.0682	$VRSC_{t-2}$	-0.0529	-0.1389	$VRSC_{t-2}$	-0.0481	0.5033
$VRSS_{t-3}$	0.0123	0.1964	$VRSC_{t-3}$	0.0094	-0.1314	$VRSC_{t-3}$	0.0544	-0.9350**
$VRSS_{t-4}$	-0.0522	-0.0045	$VRSC_{t-4}$	0.0901**	0.0623	$VRSC_{t-4}$	0.0804*	0.3228
$VRSS_{t-5}$	0.0555	0.2227	$VRSC_{t-5}$	-0.0095	0.2852	$VRSC_{t-5}$	0.0205	0.3573
$VRSS_{t-6}$	-0.0022	0.2673**	$VRSC_{t-6}$	0.0873**	0.1050	$VRSC_{t-6}$	0.0831	-0.4420
			$VRSC_{t-7}$	-0.0020	-0.0557			
			$VRSC_{t-8}$	-0.0055	-0.0606			
			$VRSC_{t-9}$	0.1127**	-0.1019			
$VRSM_{t-1}$	0.0429**	0.1510**	$VRSS_{t-1}$	-0.0034	0.2150**	$VRSM_{t-1}$	-0.00005	0.1552**
$VRSM_{t-2}$	-0.0273	0.0197	$VRSS_{t-2}$	0.0127	0.0015	$VRSM_{t-2}$	0.00007	0.0076
$VRSM_{t-3}$	-0.0116	-0.0393	$VRSS_{t-3}$	-0.0046	0.0195	$VRSM_{t-3}$	0.0068	-0.0254
$VRSM_{t-4}$	0.0143	0.0076	$VRSS_{t-4}$	-0.0119	-0.0290	$VRSM_{t-4}$	-0.0060	0.0112
$VRSM_{t-5}$	0.0201	0.0130	$VRSS_{t-5}$	-0.0094	0.0665	$VRSM_{t-5}$	-0.0013	0.0244
$VRSM_{t-6}$	0.0221	0.1210**	$VRSS_{t-6}$	-0.0027	0.0014	$VRSM_{t-6}$	-0.0006	0.1406**
			$VRSS_{t-7}$	0.0065	0.0332			
			$VRSS_{t-8}$	-0.0116	-0.0500			
			$VRSS_{t-9}$	0.0188	0.0417			
OI	-0.0063**	-0.0157**	*OI*	-0.0067**	-0.0067**	*OI*	-0.0072**	-0.0176**
VOL	0.0097**	0.0186**	*VOL*	0.0205**	0.0096**	*VOL*	0.0220**	0.0184**

**Denotes significant at the 1% level *Denotes significant at the 5% level
Volume coefficients are scaled by a factor of 1,000

Table 2: Structural Equations for Range Measure

VRS	Soybean	Mungbean	VRS	Copper	Soybean	VRS	Copper	Mungbean
Constant	4.5478**	10.1062**	*Constant*	8.2966**	11.8961**	*Constant*	4.4000**	18.6805**
$RangeC_{t-1}$	0.1423**	-0.1070	$RangeS_{t-1}$	0.1431**	-0.0346	$RangeC_{t-1}$	0.1562**	0.0151
$RangeC_{t-2}$	0.0182	0.1472*	$RangeS_{t-2}$	0.0771	-0.0103	$RangeC_{t-2}$	0.0070	0.2717**
$RangeC_{t-3}$	0.0596	-0.1162	$RangeS_{t-3}$	0.0371	0.1424**	$RangeC_{t-3}$	0.0628	-0.2025
$RangeC_{t-4}$	-0.0162	0.0359	$RangeS_{t-4}$	-0.0366	0.0248	$RangeC_{t-4}$	-0.0096	-0.0758
$RangeC_{t-5}$	-0.0225	-0.0376	$RangeS_{t-5}$	0.0187	0.1067	$RangeC_{t-5}$	-0.0032	-0.0678
$RangeC_{t-6}$	0.0453	0.0878	$RangeS_{t-6}$	-0.0191	0.0657	$RangeC_{t-6}$	0.0515	-0.0233
$RangeC_{t-7}$	0.0282	-0.0630	$RangeS_{t-7}$	0.0374	0.0889	$RangeC_{t-7}$	0.0463	-0.0803
$RangeC_{t-8}$	0.0029	0.0448	$RangeS_{t-8}$	-0.0264	0.0288	$RangeC_{t-8}$	0.0034	-0.0509
$RangeC_{t-9}$	0.0342	-0.0102	$RangeS_{t-9}$	0.0458	0.0792	$RangeC_{t-9}$	0.0713*	-0.0425
$RangeC_{t-10}$	0.1126**	-0.1483						
$RangeS_{t-1}$	0.0186	0.1666**	$RangeM_{t-1}$	0.0686**	0.1604**	$RangeM_{t-1}$	0.0154	0.1681**
$RangeS_{t-2}$	0.0259	0.0621	$RangeM_{t-2}$	-0.0155	-0.0071	$RangeM_{t-2}$	-0.0127	0.0100
$RangeS_{t-3}$	-0.0183	0.0429	$RangeM_{t-3}$	-0.0042	-0.0417	$RangeM_{t-3}$	0.0207	-0.0287
$RangeS_{t-4}$	-0.0464	-0.0313	$RangeM_{t-4}$	0.0073	-0.0430	$RangeM_{t-4}$	-0.0232	-0.0227
$RangeS_{t-5}$	0.0100	0.0224	$RangeM_{t-5}$	-0.0196	0.1093**	$RangeM_{t-5}$	0.0106	0.1139**
$RangeS_{t-6}$	0.0050	-0.0141	$RangeM_{t-6}$	-0.0074	-0.1451	$RangeM_{t-6}$	0.0012	0.0152
$RangeS_{t-7}$	-0.0054	0.0352	$RangeM_{t-7}$	0.0446	0.0622	$RangeM_{t-7}$	0.0189	0.0661
$RangeS_{t-8}$	0.0283	-0.0246	$RangeM_{t-8}$	-0.0290	-0.0449	$RangeM_{t-8}$	-0.0173	-0.0445
$RangeS_{t-9}$	-0.0020	0.0396	$RangeM_{t-9}$	0.0023	-0.0937	$RangeM_{t-9}$	0.0028	-0.0955**
$RangeS_{t-10}$	-0.0141	0.0217						
OI	-0.1271**	-0.0746**	*OI*	-0.0735**	-0.1182**	*OI*	-0.125**	-0.115**
VOL	0.4134**	0.1120**	*VOL*	0.1141**	0.1486**	*VOL*	0.415**	0.137**

**Denotes significant at the 1% level *Denotes significant at the 5% level
Volume coefficients are scaled by a factor of 1,000

Table 3. Feedback Measures: Mungbean-Soybean

According to Geweke (1982), the linear dependence (or total information flow) between two data series, series s and series, f is measured as follows: $\hat{F}_{s,f} = \ln[\hat{\sigma}^2_{us} \times \hat{\sigma}^2_{uf} / \det(\hat{\Sigma})]$. The linear feedback from series *s* to series *f* is measured by: $\hat{F}_{s\to f} = \ln(\hat{\sigma}^2_{uf} / \hat{\sigma}^2_{f})$. Similarly, the linear feedback from series *f* to series *s* is measured by: $\hat{F}_{s\leftrightarrow f} = \ln(\hat{\sigma}^2_{s} \times \hat{\sigma}^2_{f} / \det \hat{\Sigma})$. Define the instantaneous linear feedback measure between the two markets as follows: , then we have $\hat{F}_{s,f} = \hat{F}_{s\to f} + \hat{F}_{f\to s} + \hat{F}_{s\leftrightarrow f}$. Denote *M* as mungbeans, *S* as soybeans, and *C* as copper, we have the following results:

	Range	VRS		Range	VRS		Range	VRS
$\hat{F}_{M\leftrightarrow S}$	9.66389**	3.7468**	$\hat{F}_{S\leftrightarrow C}$	10.69697**	5.9437**	$\hat{F}_{M\leftrightarrow C}$	10.69697**	5.9437**
$\hat{F}_{M\to S}$	0.04536**	0.0797**	$\hat{F}_{S\to C}$	0.04817**	0.0828**	$\hat{F}_{M\to C}$	0.04327**	0.0964**
$\hat{F}_{S->M}$	0.05834**	0.0632**	$\hat{F}_{C->S}$	0.05355**	0.1022**	$\hat{F}_{C->M}$	0.04742**	0.0613**
$\hat{F}_{M,S}$	9.97676**	3.8897**	$\hat{F}_{S,C}$	10.79868**	6.1287**	$\hat{F}_{M,C}$	10.36664**	5.0691**

**Indicates significant at the 1% level

Table 4. VAR result for Range data

Variable	*RangeM*	*RangeC*	*RangeS*	*OIM*	*OIC*	*OIS*	*VOLM*	*VOLC*	*VOLS*
Constant	14.723*	5.261*	7.894*	1.709	2.824*	7.188*	-2.779	-0.212	1.078
	(2.853)[1]	(4.85)	(4.16)	(0.51)	(2.34)	(2.03)	(-0.38)	(-0.17)	(1.77)
RangeM{1}[2]	0.203*	-0.009	0.047	-0.108	0.029	-0.034	-0.095	-0.006	-0.060
	(4.01)	(-0.48)	(1.41)	(-1.82)	(1.37)	(-0.54)	(-0.73)	(-0.29)	(-0.56)
RangeM{2}	0.038	-0.012	-0.019	0.072	0.002	0.029	0.205	0.010	-0.083
	(0.75)	(-0.63)	(-0.56)	(1.20)	(0.008)	(0.46)	(1.56)	(0.48)	(-0.76)
RangeC{1}	-0.002	0.264*	-0.172	0.142	-0.007	0.152	0.601	0.101	-0.666*
	(-0.02)	(4.91)	(-1.83)	(0.86)	(-0.12)	-(-0.87)	(1.64)	(1.65)	(-2.21)
RangeC{2}	0.045	0.058	0.120	0.115	-0.068	0.044	0.007	0.008	0.070
	(0.32)	(1.10)	(1.29)	(0.70)	(-1.16)	(0.26)	(0.02)	(0.14)	(0.24)
RangeS{1}	0.044	0.021	0.213*	-0.022	0.037	-0.048	0.132	0.015	0.137
	(0.55)	(0.69)	(3.99)	(-0.24)	(1.09)	(-0.48)	(0.64)	(0.43)	(0.80)
RangeS{2}	0.015	0.010	0.125	-0.89	-0.007	0.054	-0.166	0.022	-0.206
	(0.84)	(0.34)	(2.37)	(-0.95)	(-0.23)	(0.55)	(-0.81)	(0.64)	(-1.21)
OIM[3]*{1}*	-0.049	-0.033	-0.235	1.029*	0.026	0.026	0.541*	0.019	-0.046
	(-1.07)	(-0.19)	(-0.79)	(19.48)	(1.37)	(0.46)	(4.66)	(1.01)	(-0.48)
OIM{2}	-0.022	0.008	0.239	-0.067	-0.021	-0.041	-0.355*	-0.008	0.056
	(-0.49)	(0.16)	(0.80)	(-1.27)	(-1.09)	(-0.74)	(-3.07)	(-0.40)	(0.59)
OIC{1}	-0.103	0.080	-0.472	0.134	0.903*	-0.091	-0.533	0.209*	-0.282
	(-0.77)	(0.16)	(-0.53)	(0.85)	(16.04)	(-0.55)	(-1.55)	(3.63)	(-0.99)
OIC{2}	0.045	-0.036	0.190	0.004	-0.065	-0.109	0.611	-0.114*	-0.123
	(0.35)	(-0.74)	(0.22)	(0.03)	(-1.21)	(-0.68)	(1.84)	(-2.07)	(-0.45)
OIS{1}	0.049	0.004	0.070	0.002	0.001	0.943*	0.095	0.003	0.289*
	(1.19)	(0.29)	(0.25)	(0.04)	(0.08)	(18.33)	(0.89)	(0.17)	(3.26)
OIS{2}	-0.46	-0.007	-0.078	-0.010	-0.008	-0.042	-0.101	0.002	-0.109
	(-1.15)	(-0.46)	(-0.29)	(-0.22)	(-0.47)	(-0.85)	(-0.97)	(0.12)	(-1.27)
VOLM {1}	0.048*	0.019*	-0.026	-0.002	-0.007	0.008	0.437*	0.007	-0.047
	(2.43)	(2.53)	(-0.19)	(-0.09)	(-0.87)	-(-0.32)	(8.50)	(0.84)	(-1.11)
VOLM {2}	0.016	-0.008	-0.012	-0.004	0.005	0.013	(0.63)	0.198*	-0.006
	(0.81)	(-1.02)	(-0.09)	(-0.19)	(0.51)	(3.86)	(-0.68)	0.110*	(2.61)
VOLC{1}	0.028	0.022	0.714	-0.245	0.086	0.063	(1.54)	-0.585	0.274*
	(0.21)	(0.44)	(0.81)	(-1.57)	(0.38)	(-1.70)	(4.78)	0.132	(0.47)
VOLC{2}	0.093	0.093	0.314	-0.154	0.069	0.229	-0.052	0.203*	0.579*
	(0.71)	(1.86)	(0.36)	(-0.99)	(1.23)	(1.40)	(-0.15)	(3.58)	(2.05)
VOLS{1}	-0.026	-0.006	0.152	0.002	-0.003	0.052	-0.026	-0.012	0.428*
	(-1.03)	(-0.60)	(0.91)	(0.06)	(-0.30)	(1.66)	(-0.39)	(-1.11)	(7.97)
VOLS{2}	0.015	0.005	-0.130	0.028	0.002	0.045	0.034	0.001	0.239*
	(0.59)	(0.52)	(-0.78)	(0.97)	(0.15)	(1.46)	(0.53)	(0.14)	(4.47)

1. The number in { } represents the lag value.
2. The value in parenthesis under coefficient values is the corresponding t-statistics.
3. All coefficients for the volume series are scaled by a factor of 1,000.

Table 5. VAR result for VRS data

Variable	*RangeM*	*RangeC*	*RangeS*	*OIM*	*OIC*	*OIS*	*VOLM*	*VOLC*	*VOLS*
Constant	0.199*	0.021*	0.069*	0.206	0.340*	0.585*	-0.084	-0.111	1.489
	(3.77)[1]	(3.73)	(3.73)	(0.85)	(3.90)	(2.30)	(-0.16)	(1.25)	(1.11)
VRSM{1}[2]	0.155*	-0.053	0.043*	-0.220	0.071	-0.039	-0.874	-0.045	0.024
	(3.18)	(-1.02)	(2.56)	(-0.98)	(0.88)	(-0.17)	(-1.79)	(-0.56)	(0.06)
VRSM{2}	0.014	0.063	-0.035*	-0.122	-0.048	0.170	0.355	0.014	-0.215
	(0.28)	(1.21)	(-2.06)	(-0.54)	(-0.60)	(0.72)	(0.72)	(0.17)	(-0.52)
VRSC{1}	-0.732	0.163*	-0.293	0.165	-0.284	0.293	12.743*	-0.981	-3.085
	(-1.57)	(3.27)	(-1.80)	(0.77)	(-0.37)	(0.13)	(2.73)	(-1.25)	(-0.79)
VRSC{2}	0.017	-0.016	-0.129	0.200	-0.770	2.526	1.586	1.336	1.346
	(0.04)	(-0.35)	(-0.85)	(1.00)	(-1.08)	(1.21)	(0.37)	(1.84)	(0.37)
VRSS{1}	0.002	0.014	0.209*	-0.300	-0.042	-0.323	1.872	-0.034	-0.95
	(0.01)	(0.09)	(4.15)	(-0.45)	(-0.18)	(-0.46)	(1.29)	(-0.14)	(-0.78)
VRSS{2}	-0.037	-0.27	0.040	-0.468	0.085	-0.436	0.256	0.027	-0.863
	(-0.26)	(-0.18)	(0.81)	(-0.71)	(0.36)	(-0.63)	(0.18)	(0.11)	(-0.72)
OIM[3]*{1}*	-0.080	-0.001	0.019	1.031*	0.023	0.025	0.539*	0.020	-0.042
	(-0.70)	(-0.09)	(0.48)	(19.59)	(1.21)	(0.44)	(4.69)	(1.03)	(-0.43)
OIM{2}	-0.080	-0.004	-0024	-0.070	-0.022	-0.037	-0.367	-0.011	0.068
	(-0.70)	(-0.31)	(-0.60)	(1.34)	(-1.16)	(-0.67)	(-3.21)	(0.56)	(0.71)
OIC{1}	-0.374	-0.031	-0.030	0.133	0.906*	-0.062	-0.458	0.188*	-0.233
	(-1.10)	(-0.85)	(-0.25)	(0.85)	(16.14)	(-0.38)	(-1.35)	(3.30)	(-0.82)
OIC{2}	0.290	-0.014	-0.037	0.005	-0.065	-0.098	0.598	-0.103	-0.115
	(0.89)	(-0.41)	(-0.32)	(0.03)	(-1.20)	(-0.62)	(1.83)	(-1.87)	(-0.42)
OIS{1}	0.034	-0.001	-0.005	0.004	-0.002	0.944*	0.096	-0.005	0.292
	(0.32)	(-0.01)	(-0.12)	-(-0.01)	(-0.09)	(18.56)	(0.91)	(-0.29)	(3.30)
OIS{2}	-0.044	0.006	0.001	-0.006	-0.006	-0.051	-0.099	0.006	-0.110
	(-0.43)	(0.54)	(0.04)	(-0.12)	(-0.36)	(-1.04)	(-0.96)	(0.35)	(-1.28)
VOLM{1}	0.034	0.011*	-0.026	-0.015	-0.004	-0.013	0.440*	0.005	-0.058
	(0.70)	(2.17)	(-1.54)	(-0.66)	(-0.53)	(-0.54)	(9.05)	(0.64)	(-1.43)
VOLM{2}	0.116*	-0.005	0.018	0.010	0.006	0.012	0.223*	-0.001	0.096*
	(2.40)	(-0.88)	(1.05)	(0.46)	(0.79)	(-0.53)	(4.59)	(-0.16)	(2.37)
VOLC{1}	0.208	0.090*	0.058	-0.220	0.087	-0.015	-0.629*	0.350*	-0.100
	(0.66)	(2.69)	(0.53)	(-1.52)	(1.69)	(0.10)	(-2.00)	(6.64)	(-0.38)
VOLC{2}	0.018	0.069*	0.094	-0.177	0.060	0.177	-0.159	0.178*	0.553*
	(0.57)	(2.02)	(0.85)	(-1.21)	(1.15)	(1.16)	(-0.50)	(3.34)	(2.08)
VOLS{1}	-0.025	-0.017*	0.015	-0.000	0.004	0.049	-0.024	-0.009	0.461*
	(-0.43)	(-2.68)	(0.73)	(-0.01)	(0.39)	(1.72)	(-0.41)	(-0.92)	(9.28)
VOLS{2}	0.017	0.008	-0.012	0.030	-0.003	0.061*	0.027	0.005	0.209*
	(0.27)	(1.23)	(-0.59)	(1.09)	(-0.33)	(2.11)	(0.44)	(0.47)	(4.15)

1. The number in { } represents the lag value.
2. The value in parenthesis under coefficient values is the corresponding t-statistics.
3. All coefficients for the volume series are scaled by a factor of 10,000.

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An Examination of Higher Moments of Latin American Equity Returns Distributions

Terrance Grieb

Associate Professor of Finance
University of Idaho
College of Business and Economics
P.O. Box 443161
Moscow, ID
83844-3161
(208) 885-7140 (office)
tgrieb@uidaho.edu

Abstract

This study examines the higher-moments of the returns distribution for Latin American equity indexes. Evidence is provided that kurtosis presents more of a risk than skewness in these markets, and while non-normal distributions are observed more frequently during the middle of the year, significant differences in the degree of skewness and kurtosis across the months of the year are sporadic. Diversifying portfolio holdings across the region tends to reduce variance and kurtosis risk, but causes an increased exposure to negative skewness. Finally, while regional diversification can reduce total risk, it does little to affect systematic risk exposure levels.

I. Introduction

Investors typically focus on co-variance risk when measuring the risk-return characteristics of their portfolios, although investors should also consider dimensions of risk represented by the higher-moments of the returns distribution. Skewness (the third moment above the mean) is the degree of asymmetry present in an otherwise normal distribution. It is hypothesized that investors will accept lower returns for positive skewness and will require higher returns for negative skewness. The opposite is true for kurtosis (the fourth moment above the mean), implying that greater degrees of kurtosis should result in higher expected returns for investors. (See Tang and Shum (2003) for discussion of investor preferences and higher moments of the returns distribution.)

Early work in this area by Kraus and Litzenberger (1976), and Simkowitz and Beedles (1978) established that skewness risk has a diversifiable component and that co-skewness risk was significantly priced in U.S. securities. Friend and Westerfield (1980) established that risk factors for co-skewness risk are not constant over time.

Cremers, Kritzman, and Page (2005) examine optimal portfolio allocations under the assumptions of bilinear or S-shaped utility curves. Citing Kahneman and Tversky (1979), they argue that investors have conditional utility preferences based on market conditions (i.e., certain loss or certain gain) that are not captured by mean-variance analysis. By including skewness and kurtosis in their optimization algorithm they maximize the success rates for investors with S-shaped utility functions. They also show that kurtosis and negative skewness are less problematic for S-shaped utility functions because extreme results are not severely penalized.

Harvey and Siddique (1999) employ a GARCH-M model to observe conditional skewness pricing for U.S. equity returns. They demonstrate that conditional skewness is significantly priced and that the relationship between the three moments of the distribution is linked to seasonal variations in the conditional moments. Likewise, Chiao, Hung, and Srivasta (2003) apply a conditional four-moment CAPM to the Taiwan equity market and find that co-skewness and co-kurtosis risk are significantly priced only in up-markets for Taiwan stocks.

In contrast, Prakesh, Moncarz, and Anderson (2001) argue that the probability distributions for equities and indexes in Latin America are not Gaussian, and fit cross-sectional two-moment and four-moment CAPM frameworks to stocks from four countries to test the intertemporal stability of the parameters.[1] They provide evidence of stability for co-variance, co-skewness, and co-kurtosis risk. They also show that co-variance risk is stable for the time period observed even without co-skewness and co-kurtosis risk in the model.

Tang (1998) examines monthly patterns and diversification characteristics for sector indexes in the Hong Kong market and finds a conditional month-of-the-year effect in the higher moments of the indexes. His results indicate that diversifying across the indexes reduces the standard deviation and kurtosis of a portfolio, but that negative skewness tends to become even more negatively skewed with diversification.

This study extends the results of Prakesh, Moncarz, and Anderson (2001) by examining the month-of-the-year effects and diversification effects on the higher-moments of the returns distributions for index level returns in seven Latin American countries. The next section describes the data set and the methodology employed. Section III presents the empirical results, and Section IV concludes.

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[1] They observe a total of 28 stocks from Argentina, Brazil, Chile and Mexico, using weekly returns for 1997-1999, and they use the home country index and the 90-day LIBOR as the market and risk-free proxies.

II. Data and Methodology

The data consists of daily returns for Argentina, Brazil, Chile, Columbia, Mexico, Peru and Venezuela for the five year period from June 1, 2000 to May 31, 2005, calculated as the natural log of the price relative and denominated in U.S. dollars.[2] The data is obtained from the Morgan Stanley Capital International, Inc. (MSCI) database. To form the country indexes all listed securities in the market are free-float adjusted and screened by size and liquidity, then a value-weighted portfolio is formed that captures 85% of the market capitalization. This methodology captures the economic makeup of each market and reflects the value available to foreign investors with minimal liquidity constraints. Where needed, the MSCI World and U.S. indexes are used to proxy the overall market returns.

The expected values of skewness (denoted g_1) and kurtosis (denoted g_2) for a normal distribution are zero and three. The skewness and kurtosis values are estimated for each country's returns as follows:

$$g_1 = \left(\frac{n}{(n-1)(n-2)}\right)\sum_{i=1}^{n}\frac{(x_i-\bar{x})^3}{s^3} \tag{1}$$

and

$$g_2 = \left(\frac{n(n+1)\sum_{i=1}^{n}(x_i-\bar{x})^4}{(n-1)(n-2)(n-3)s^4}\right) - \left(\frac{3(n-1)^2}{(n-2)(n-3)}\right) \tag{2}$$

where n is the number of observations for country i, and s is the sample standard deviation of the returns. The estimate of kurtosis in equation (2) is adjusted so that it has an expected value of zero. Accordingly, both g_1 and g_2 are normally distributed with an expected value of zero and a standard deviation of $\sqrt{6/n}$ for g_1 and $\sqrt{24/n}$ for g_2. The Jarque-Bera (J-B) statistic is used to test for normality of the distributions. The J-B statistic takes the form:

$$JB = \frac{n}{6}\left(g_1^2 + \frac{(g_2^2)}{4}\right) \tag{3}$$

and is distributed as a χ^2 with 2 degrees of freedom.

Consistent with a procedure developed by Tang (1998), the equality of skewness and kurtosis between two samples is tested by standardizing the returns and applying a Kolmogorov-Smirnov non-parametric test (K-S test) for equal distributions. The K-S test compares the maximum relative distance between the empirical cumulative distribution function (ECDF) of two samples (denoted F1 and F2). The D-statistic is estimated as:

$$D = \max_{1\le i\le n}\left|F1_i - F2_i\right|,\ \forall_{i=1\ to\ n} \tag{4}$$

and the approximate critical value is given by:

$$D_a = c(\alpha)\sqrt{\frac{n_1+n_2}{n_1 n_2}} \tag{5}$$

[2] Returns for Argentina for January 7 and January 17, 2002, and for Venezuela for November 28, 2003 are excluded from the dataset. The Mercado de Valores de Buenos Aries was closed from January 8 – 16, 2002 due to the economic crisis surrounding the devaluation of the Argentine peso. The resulting index returns for the days before and after the closure were -33.6% and -23.9%, respectively. On November 28, 2003, MSCI switched its standard spot rate for the Venezuela bolivar to a notional rate, generating a dollar-denominated index return of -47.8% for that date.

where c(α) equals 1.63 for $\alpha = 1\%$ and 1.36 for $\alpha = 5\%$, and the values n_1 and n_2 are the sample sizes for the F1 and F2 distributions. Since standardizing returns generates a mean of zero and a standard deviation of one without altering skewness or kurtosis, the K-S test applied to standardized returns tests the null hypothesis of equality in the higher moments of the two sample distributions.

Finally, the co-variance risk (β_j), co-skewness risk (γ_j), and co-kurtosis risk (δ_j) for each country relative to an overall market is estimated as follows:

$$\beta_j = \frac{\sum_{t=1}^{n}\left[\left(R_{mt}-\bar{R}_m\right)\left(R_{jt}-\bar{R}_j\right)\right]}{\sum_{t=1}^{n}\left(R_{mt}-\bar{R}_m\right)^2}, \tag{6}$$

$$\gamma_j = \frac{\sum_{t=1}^{n}\left[\left(R_{mt}-\bar{R}_m\right)^2\left(R_{jt}-\bar{R}_j\right)\right],}{\sum_{t=1}^{n}\left(R_{mt}-\bar{R}_m\right)^3} \tag{7}$$

and,

$$\delta_j = \frac{\sum_{t=1}^{n}\left[\left(R_{mt}-\bar{R}_m\right)^3\left(R_{jt}-\bar{R}_j\right)\right]}{\sum_{t=1}^{n}\left(R_{mt}-\bar{R}_m\right)^4}, \tag{8}$$

where R_{jt} and R_{mt} are the time period t returns for country j and the overall market, respectively.

III. Empirical Results

Monthly skewness and kurtosis effects by country

To test for seasonal patterns in the higher moments of the returns distribution the data is sorted by month for each country and tested for skewness, kurtosis, and normality. Table 1 shows that significant levels of skewness and kurtosis occur throughout the year, as do significant violations of normality. The general result is that normality cannot be assumed in the returns distributions for any of the countries. However, two interesting patterns exist. First, the violations of normality are more widespread in the middle of the year (May through October) than the turn of the year (November through March). Of particular note is September when all seven countries exhibit non-normal distributions and significant degrees of both skewness and kurtosis. Second, significant kurtosis occurs more frequently than significant skewness, which implies that controlling for extreme-return events should be of more concern for investors than controlling for a bias towards positive or negative returns.

Next, the Kolmogorov-Smirnov test is applied to test the null hypothesis of equality in returns distributions for each country across the months of the year. Comparing all combinations of months results in 66 tests per country. Table 2 reports only the months where a significant difference in the returns distribution is observed. In general, there appear to be very few significant differences. Brazil, Chile, Colombia, and Mexico all show three or fewer combinations that reject the null hypothesis. Peru shows only six combinations, although five of those involve the month of April, indicating a small monthly effect. More pronounced monthly effects are observed for Argentina and Venezuela.[3] In the case of Argentina, all observed rejections of the null involve either December or January. In almost all of these combinations the month at the

[3]Excluding 1/7/02 and 1/17/02 for Argentina, and 11/28/03 for Venezuela should actually bias the data towards a failure to reject the null hypothesis. The results obtained above support the observations that the monthly effects observed are not driven by a few outlier events. As a robustness check the tests were run including those dates and the results were similiar.

Table 1
Skewness, Kurtosis, and Normality Measures by Month

Panel A: Skewness

	Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec
Argentina	0.83a	-1.71a	-0.89	-0.244	-0.36	0.54b	-0.33	0.13	-0.71a	-0.14	-0.18	1.82a
Brazil	-0.13	-0.27	-0.42c	-0.258	-0.40c	-0.36	-0.64a	1.36a	-0.87a	0.56b	0.18	0.04b
Chile	-0.52	0.18	-0.39	0.132	-0.26	-0.30	-0.16	0.07	-1.57a	0.03	-0.38	-0.22
Colombia	0.27b	0.12	-0.43	0.669a	0.41	-0.49b	0.22	0.85a	0.75a	0.24	0.27	0.38
Mexico	0.23	0.03	-0.39	0.097	0.59b	-0.01	0.00	-0.14	-0.99a	0.16	0.03	-0.42
Peru	-1.43a	0.40	0.61b	-1.482a	-0.03	0.59b	-0.49b	-0.01	-1.79a	-0.26	0.51b	-0.96a
Venezuela	1.83a	1.03a	0.42	-0.123	0.61b	-0.30	-0.81a	-0.22	1.05a	-0.14	0.03	-0.02

Panel B: Kurtosis

	Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec
Argentina	4.09a	9.63a	12.73a	0.11	0.67	1.92a	4.45a	2.04a	1.33a	2.79a	0.861	7.96a
Brazil	1.85a	0.12	0.29	0.55	1.12b	1.01b	0.14	7.05a	4.28a	0.54	0.697	0.52
Chile	-0.18	0.43	-0.12	-0.25	0.49	0.49b	1.04	0.43	4.50a	-0.15	0.750	0.57
Colombia	0.23	1.25b	1.56a	1.23b	5.47a	0.85	1.50a	3.99a	2.15a	0.20	1.286a	2.69a
Mexico	0.40	0.93	0.48	-0.31	5.02a	2.14a	5.18a	1.95a	4.20a	0.37	1.171b	7.84a
Peru	12.26a	0.84	0.41	4.34a	2.61a	1.33a	3.73a	1.88a	6.48a	1.05b	2.575a	7.96a
Venezuela	11.36a	5.86a	1.67a	5.16a	1.55a	1.22b	2.24a	2.05a	5.40a	2.57a	1.339b	1.16

Panel C: Jarque-Bera test for normality

	Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec
Argentina	80.46a	426.5a	743.8a	1.03	4.28	21.34a	92.59a	19.38a	16.45a	36.58a	3.88	318.9a
Brazil	15.75a	1.22	3.58	2.50	8.70b	6.71b	7.69b	264.1a	94.24a	7.11b	2.77	1.20
Chile	5.05	1.29	2.84	0.58	2.37	2.61	5.48	0.95	129.5a	0.12	5.07	2.23
Colombia	1.55	6.70b	14.29a	14.49a	141.2a	7.39b	11.18a	86.29a	30.41a	1.24	8.67b	34.52a
Mexico	1.72	3.61	3.77	0.58	123.0a	20.22a	123.9a	18.02a	95.33a	1.11	6.13b	274.4a
Peru	713.2a	5.52	7.49b	119.7a	31.39a	13.83a	67.31a	16.40a	242.1a	5.99b	33.81a	296.0a
Venezuela	623.0a	156.0a	15.25a	114.5a	17.52a	7.75b	34.23a	20.06a	148.3a	30.78a	7.79b	5.85

a = significant at the 1% level.
b = significant at the 5% level.
c = significant at the 10% level

turn of the year exhibits positive skewness and a large degree of kurtosis, whereas the comparison month has negative skewness and a smaller degree of kurtosis. For Venezuela, all combinations involving November are significant. November has a low level of skewness and kurtosis relative to most of the other months, but no pronounced pattern is distinguishable.

Considering Tables 1 and 2 together, the results suggest that while kurtosis may be more prevalent than skewness across the region, and that both positive and negative skewness events are possible, the only pronounced month-of-the-year effect in returns distributions are for Argentina, which exhibits a December and January effect, and Venezuela, which exhibits a November effect. Some evidence of an April effect is found for Peru.

Portfolio effects

To determine the effect of regional diversification, portfolio returns are observed as the number of indexes in the portfolio is increased. Returns for equal-weighted portfolios consisting of one through seven countries are estimated, and all possible combinations of indexes are included.[4] This approach generates the most portfolio observations and also prevents the results of any one country from having an undue influence on the portfolio characteristics.

Table 3 shows the results of increasing regional exposure on average levels of standard deviation, skewness, and kurtosis. Standard deviation decreases uniformly for all months as

[4] The possible combinations for portfolios of size one through seven are 7, 21, 35, 35, 21, 7, and 1, respectively.

Table 2
Months with Significantly Different Returns Distributions by Country

Argentina		Brazil		Chile		Colombia		Mexico		Peru		Venezuela	
Months	K-S	Months	K-S	Months	K-S	Months	K-S	Months	K-S	Months	K-S	Months	K-S
Jan-Feb	0.105 c	Jul-Aug	0.162 c	--	--	Mar-Oct	0.170c	Sep-Oct	0.175c	Mar-Apr	0.190 b	Jan-Nov	0.237 a
Jan-Mar	0.173 c	Sep-Oct	0.188 b			May-Jun	0.166 c			Apr-June	0.171 c	Feb-Nov	0.278 a
Jan-Apr	0.222 b					May-Oct	0.186 b			Apr-Aug	0.174 c	Mar-Jul	0.171 c
Jan-May	0.193 b									Apr-Nov	0.182 c	Mar-Nov	0.332 a
Jan-Jun	0.221 b									Apr-Dec	0.201 b	Mar-Dec	0.171 c
Jan-Jul	0.240 a									Sep-Dec	0.170 c	Apr-Nov	0.216 b
Jan-Aug	0.186 b											May-Jul	0.167 c
Jan-Sep	0.211 b											May-Nov	0.301 a
Jan-Oct	0.212 b											Jun-Nov	0.256 a
Jan-Nov	0.236 a											Jul-Nov	0.200 b
Jan-Dec	0.248 a											Aug-Nov	0.303 a
Feb-Dec	0.177 c											Sep-Nov	0.281 a
May-Dec	0.186 b											Oct-Nov	0.290 a
Jul-Dec	0.187 b												
Nov-Dec	0.228 a												
Aug-Dec	0.166 c												
Sep-Dec	0.182 c												
Nov-Dec	0.169 c												

a = significant at the 1% level. b = significant at the 5% level. c = significant at the 10% level.

portfolio scope increases, showing that diversification decreases unsystematic risk. In contrast, the results for skewness strongly indicate that an increase in diversification causes the distribution of returns to become more negatively skewed, which is an undesirable result for risk-averse investors. Similar to the results for standard deviation, the degree of kurtosis shows a tendency to decrease as portfolio scope increases, which is consistent with an improvement in the risk profile for the fourth moment as the portfolio is diversified across the region.

The results of Table 3 indicate that as a portfolio is diversified across Latin American markets the level of exposure to variance and kurtosis risk is decreased, but at a cost of increased exposure to negative skewness. Given that kurtosis seems to be a larger factor in the non-normal characteristics of Latin American returns than skewness, it seems that regional diversification presents a worthwhile tradeoff for risk-averse investors.

To gain more understanding of the portfolio risk characteristics for a diversified global investor the average levels of co-variance, co-skewness, and co-kurtosis are observed for portfolios with increasing levels of regional diversification across the months of the year. Table 4 shows the results of the systematic risk measures for moments two through four using the MSCI World index as the market proxy.[5] The results generally show that regional diversification does not noticeably change the degree of systematic risk for those moments. An important exception is a decrease in co-skewness risk for July and September, which is of particular interest given that previous results indicate an increased exposure to negative skewness in these months.

The overall results imply that diversification can reduce unsystematic risk in the second and fourth moments. Diversification has little effect on the systematic risk of portfolios that are already internationally diversified, with the exception of co-skewness risk for certain months.

Tang (1998) argues that if the correlation between indexes is higher for negative returns than for positive returns – indicating that bad news is more quickly reflected in returns than positive news – then negative skewness would be induced in a diversified portfolio. Table 5 observes correlations between positive and negative returns to test for this effect. The null hypothesis is rejected only 14 times. However, in every significant case the t-score is negative, indicating that bad news was processed more quickly than good news. In addition, the results

[5] The same test is run using the MSCI U.S. index as the market portfolio. The results are not shown here to save space, but are very similar to the results reported.

Table 3
Diversification Effect on Standard Deviation, Skewness, and Kurtosis

Month	1	2	3	4	5	6	7
AVERAGE STANDARD DEVIATION BY PORTFOLIO SCOPE (Number of Indexes in Portfolio)							
Jan	2.005	1.569	1.364	1.253	1.179	1.122	1.081
Feb	1.876	1.453	1.241	1.116	1.035	0.972	0.924
Mar	1.782	1.407	1.238	1.152	1.086	1.045	1.014
Apr	1.764	1.412	1.249	1.162	1.107	1.064	1.035
May	1.843	1.513	1.361	1.286	1.245	1.210	1.185
Jun	1.671	1.289	1.125	1.041	0.990	0.944	0.914
Jul	1.679	1.277	1.099	1.008	0.940	0.889	0.854
Aug	1.527	1.182	1.026	0.953	0.889	0.846	0.816
Sep	1.636	1.307	1.171	1.110	1.067	1.032	1.011
Oct	1.676	1.316	1.162	1.091	1.045	1.002	0.976
Nov	1.917	1.497	1.272	1.145	1.080	1.012	0.957
Dec	1.621	1.268	1.107	1.016	0.951	0.904	0.870
AVERAGE CO-SKEWNESS RISK BY PORTFOLIO SCOPE							
Jan	0.155	0.221	0.137	0.024	-0.068	-0.152	-0.228
Feb	-0.033	-0.047	-0.094	-0.118	-0.114	-0.134	-0.148
Mar	-0.213	-0.219	-0.254	-0.290	-0.322	-0.339	-0.357
Apr	-0.173	-0.314	-0.325	-0.317	-0.305	-0.288	-0.270
May	0.081	-0.198	-0.400	-0.548	-0.666	-0.762	-0.833
Jun	-0.047	-0.063	-0.085	-0.111	-0.130	-0.136	-0.129
Jul	-0.313	-0.419	-0.494	-0.537	-0.558	-0.602	-0.636
Aug	0.292	0.182	0.070	0.019	-0.075	-0.149	-0.211
Sep	-0.590	-0.822	-1.072	-1.252	-1.403	-1.497	-1.588
Oct	0.064	0.006	-0.041	-0.076	-0.142	-0.189	-0.220
Nov	-1.130	-2.092	-2.525	-2.686	-3.002	-2.929	-2.736
Dec	0.090	0.242	0.275	0.267	0.235	0.199	0.169
AVERAGE CO-KURTOSIS RISK BY PORTFOLIO SCOPE							
Jan	4.288	3.199	2.287	1.898	1.794	1.930	2.212
Feb	2.721	2.511	1.960	1.320	0.817	0.514	0.320
Mar	2.432	2.128	1.611	1.150	0.904	0.731	0.652
Apr	1.547	1.254	0.839	0.478	0.162	-0.003	-0.095
May	2.418	1.831	1.793	2.007	2.253	2.551	2.791
Jun	1.282	1.185	1.206	1.222	1.354	1.491	1.647
Jul	2.609	1.555	1.091	0.785	0.576	0.466	0.379
Aug	2.770	2.398	2.232	2.228	2.123	2.168	2.241
Sep	4.048	3.199	3.544	4.106	4.747	5.183	5.624
Oct	1.048	0.671	0.384	0.192	0.025	-0.075	-0.158
Nov	12.136	18.422	20.475	20.254	21.157	19.243	16.624
Dec	4.098	2.643	1.836	1.263	0.867	0.569	0.377

show that the majority of the tests resulted in negative t-scores. The highest concentration of negative t-scores occurs in the middle of the year (and in September in particular), which is consistent with the normality violations observed in Table 1. Overall, there is some evidence that a differential reaction time for bad news vs. good news imposes some negative bias on skewness.

Table 4
Diversification Effect on Systematic Risk Exposure to the World Market Index

AVERAGE COVARIANCE RISK BY PORTFOLIO SCOPE							
	Number of Indexes in Portfolio						
Month	1	2	3	4	5	6	7
Jan	0.620	0.619	0.636	0.637	0.622	0.617	0.616
Feb	0.383	0.385	0.394	0.394	0.392	0.392	0.393
Mar	0.487	0.489	0.495	0.495	0.495	0.494	0.495
Apr	0.543	0.543	0.552	0.553	0.549	0.545	0.546
May	0.862	0.867	0.863	0.864	0.876	0.874	0.873
Jun	0.702	0.708	0.724	0.727	0.733	0.729	0.734
Jul	0.303	0.304	0.309	0.306	0.315	0.313	0.315
Aug	0.449	0.449	0.457	0.452	0.451	0.449	0.449
Sep	0.462	0.462	0.470	0.466	0.463	0.460	0.460
Oct	0.483	0.483	0.482	0.478	0.489	0.483	0.483
Nov	0.461	0.462	0.466	0.462	0.469	0.466	0.467
Dec	0.442	0.443	0.455	0.454	0.447	0.444	0.443
AVERAGE CO-SKEWNESS RISK BY PORTFOLIO SCOPE							
Jan	5.994	4.825	3.551	3.158	3.207	3.238	3.614
Feb	3.197	3.146	3.173	3.134	3.027	2.977	2.950
Mar	2.853	2.861	2.797	2.799	2.896	2.893	2.901
Apr	0.289	0.287	0.287	0.285	0.285	0.280	0.278
May	4.802	4.827	4.821	4.863	4.901	4.941	4.972
Jun	0.803	0.815	0.843	0.844	0.872	0.876	0.893
Jul	-1.732	-1.644	-1.545	-1.454	-1.274	-1.166	-1.044
Aug	0.037	0.036	0.036	0.020	0.040	0.029	0.027
Sep	0.953	0.874	0.913	0.952	0.707	0.568	0.494
Oct	0.158	0.156	0.146	0.133	0.150	0.148	0.146
Nov	-0.593	-0.585	-0.468	-0.436	-0.590	-0.545	-0.533
Dec	-1.022	-1.071	-1.075	-1.113	-1.192	-1.185	-1.192
AVERAGE CO-KURTOSIS RISK BY PORTFOLIO SCOPE							
Jan	0.477	0.474	0.501	0.501	0.469	0.464	0.461
Feb	0.280	0.282	0.293	0.294	0.293	0.294	0.297
Mar	0.488	0.489	0.492	0.491	0.492	0.490	0.490
Apr	0.513	0.513	0.519	0.519	0.519	0.515	0.516
May	0.799	0.801	0.795	0.793	0.805	0.805	0.804
Jun	0.716	0.723	0.743	0.745	0.752	0.750	0.757
Jul	0.273	0.275	0.275	0.269	0.286	0.283	0.285
Aug	0.548	0.549	0.554	0.549	0.553	0.550	0.550
Sep	0.425	0.424	0.432	0.425	0.424	0.422	0.421
Oct	0.340	0.340	0.337	0.330	0.343	0.339	0.339
Nov	0.431	0.434	0.440	0.436	0.447	0.446	0.449
Dec	0.602	0.603	0.614	0.610	0.608	0.606	0.606

Conditional skewness and kurtosis

To examine the degree of skewness and kurtosis in the distributions based on bull-market vs. bear-market conditions the sample is partitioned using an October 1, 2002 split date. Poor results across the region were evident prior to this split date, and all seven countries demonstrated consistently positive returns during the post-split period. This split date is consistent with bull- vs. bear-market conditions in the global market, indicating that these results are based on market prices rather than market structure factors.

Table 5
Correlation Between Indexes for Positive and Negative Returns

		Argentina	Brazil	Chile	Colombia	Mexico	Peru	Venezuela
Jan	Positive	0.311	0.185	0.238	0.014	0.308	0.176	-0.071
	Negative	0.016	0.409	0.323	-0.063	0.239	0.313	0.147
	t-stat	1.480	-1.246	-0.463	0.383	0.370	-0.725	-1.069
Feb	Positive	-0.110	0.091	0.203	0.122	0.352	0.288	-0.137
	Negative	0.141	0.033	0.400	0.034	0.119	0.116	-0.297
	t-stat	-1.206	0.279	-1.048	0.427	1.204	0.866	0.798
Mar	Positive	0.060	0.298	0.410	0.072	0.157	0.129	-0.147
	Negative	0.169	0.462	0.198	0.065	0.436	0.223	0.023
	t-stat	-0.553	-0.974	1.190	0.035	-1.562	-0.492	-0.831
Apr	Positive	0.267	0.536	0.332	0.042	0.245	0.178	-0.139
	Negative	-0.001	0.543	0.424	0.211	0.359	0.205	0.193
	t-stat	1.325	-0.051	-0.535	-0.844	-0.625	-0.138	-1.631**c**
May	Positive	0.121	0.449	0.578	-0.037	0.215	0.298	0.053
	Negative	0.416	0.670	0.525	0.230	0.470	0.422	0.473
	t-stat	-1.613**c**	-1.665**c**	0.390	-1.382	-1.475	-0.707	-2.325**b**
June	Positive	0.199	0.201	0.240	0.038	0.452	0.157	0.016
	Negative	0.246	0.385	0.377	-0.217	0.371	0.052	0.343
	t-stat	-0.241	-1.005	-0.748	1.284	0.486	0.529	-1.651**c**
July	Positive	-0.160	0.222	0.227	0.158	0.122	-0.012	-0.214
	Negative	-0.016	0.354	0.297	0.020	0.372	0.227	0.107
	t-stat	-0.741	-0.735	-0.389	0.707	-1.377	-1.216	-1.640**c**
Aug	Positive	-0.160	0.185	0.324	-0.090	0.327	0.229	0.165
	Negative	0.208	0.549	0.444	0.287	0.032	0.309	0.163
	t-stat	-1.890**c**	-2.196**b**	-0.710	-1.962**c**	1.566	-0.426	0.009
Sept	Positive	0.232	-0.030	-0.063	0.197	0.317	-0.088	0.080
	Negative	0.406	0.539	0.574	0.267	0.603	0.363	0.367
	t-stat	-0.967	-3.157**a**	-3.516**a**	-0.374	1.842**c**	-2.334**b**	-1.515
Oct	Positive	0.285	0.192	0.451	0.017	0.231	0.131	0.203
	Negative	0.352	0.434	0.391	-0.263	0.162	0.072	0.222
	t-stat	-0.381	-1.391	0.376	1.469	0.369	0.303	-0.100
Nov	Positive	0.069	0.242	-0.253	-0.008	0.318	0.101	0.233
	Negative	0.214	0.377	0.237	-0.015	0.308	0.104	-0.055
	t-stat	-0.742	-0.742	-2.485**b**	0.034	0.056	-0.011	1.438
Dec	Positive	0.015	0.267	0.027	0.183	0.282	0.093	0.285
	Negative	0.061	-0.022	0.032	0.304	0.127	0.294	0.135
	t-stat	-0.223	1.410	-0.022	-0.641	0.809	-1.045	0.776

Notes: "Positive" indicates that the correlations in that row are between indexes for days with positive returns, "Negative" indicates correlations for negative return days.
a = significant at the 1% level, b = significant at the 5% level, c = significant at the 10% level.

Table 6 examines the level of both skewness and kurtosis for each country in both the pre-split and post-split periods. Overall there was no persuasive evidence of a shift in skewness. Kurtosis remained stable for most countries, although Argentina, Mexico and Venezuela all saw a notable decrease in kurtosis in the post-split period. These results are consistent with the stability of the of higher-moment risk factors documented by Prakash, Moncarz, and Anderson (2001). This implies that risk exposure to higher-moments of the distribution is less conditional on up-markets vs. down-markets for Latin American indexes than for other regions.

Table 6
Skewness and Kurtosis Comparison for Bull vs. Bear Markets

Skewness comparison for bull vs. bear market

		Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec
Argentina	Bear Market	0.432	-1.456	-0.690	-0.232	0.084	0.982	-0.184	0.212	-0.768	0.086	-0.174	1.164
	Bull Market	0.178	-0.726	-0.001	-0.151	-0.688	-0.831	-0.263	-0.354	-0.096	0.424	0.253	2.057
Brazil	Bear Market	0.887	-0.203	-0.655	-0.401	-0.250	-0.472	-0.348	1.475	-0.450	0.343	0.521	0.175
	Bull Market	-0.394	-0.327	-0.194	-0.159	-0.545	0.608	-0.224	-0.805	-0.252	0.609	-0.102	-0.185
Chile	Bear Market	-0.936	0.212	0.228	0.435	0.174	-0.550	-0.231	0.226	-1.299	0.126	0.206	0.097
	Bull Market	-0.305	0.118	-0.599	-0.018	-0.296	-0.192	0.575	-0.149	0.077	0.013	-0.829	-0.398
Colombia	Bear Market	0.619	0.256	-0.082	0.253	2.121	-0.425	0.562	1.027	0.121	0.342	1.074	0.651
	Bull Market	0.216	0.004	-0.562	0.703	0.099	-0.537	-0.077	0.696	0.766	0.229	-0.126	0.254
Mexico	Bear Market	0.320	-0.243	-0.754	0.224	2.027	0.075	0.133	-0.113	-0.548	0.095	-0.344	-0.389
	Bull Market	-0.280	0.261	-0.020	-0.053	-0.757	-0.333	-0.144	-0.721	-1.101	0.263	0.573	-0.407
Peru	Bear Market	0.299	0.427	0.301	-2.238	0.255	0.602	-1.192	-0.269	-1.967	-0.663	-0.738	0.277
	Bull Market	-1.188	0.384	0.719	-1.250	-0.100	0.452	0.776	0.242	-0.280	-0.393	0.452	-0.928
Venezuela	Bear Market	0.459	0.004	0.205	0.250	0.775	-0.352	-0.425	-0.127	1.600	-1.225	0.007	0.266
	Bull Market	2.039	1.626	0.594	-0.736	0.413	0.047	-1.364	-0.395	-0.124	0.469	0.068	0.008

Kurtosis comparison for bull vs. bear market

		Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec
Argentina	Bear Market	2.265	7.305	6.937	-1.035	-0.463	1.913	2.690	1.587	1.253	0.958	0.156	2.958
	Bull Market	-0.225	1.608	-0.133	1.333	1.990	1.320	-0.217	0.791	-0.357	0.160	0.738	10.310
Brazil	Bear Market	1.828	0.636	1.006	1.255	-0.226	0.570	-0.601	5.455	2.749	-1.151	1.111	0.840
	Bull Market	1.452	-0.102	-0.597	-0.051	0.956	0.008	-0.207	0.535	-0.001	1.029	0.266	-0.477
Chile	Bear Market	0.979	-0.249	-0.494	-0.215	-0.252	0.390	0.905	0.554	2.839	-0.222	0.554	0.548
	Bull Market	-0.613	0.932	-0.197	-0.331	-0.199	0.437	-0.055	0.065	0.006	-0.080	0.827	0.710
Colombia	Bear Market	-0.114	2.643	0.302	-0.383	7.121	0.154	0.876	5.979	-0.037	0.706	1.756	0.061
	Bull Market	0.200	0.822	1.357	1.245	4.513	2.580	1.633	1.312	1.697	0.028	1.279	2.958
Mexico	Bear Market	-0.322	1.845	0.437	-0.514	8.967	1.397	3.321	1.079	2.460	-0.456	0.120	5.236
	Bull Market	0.434	0.083	0.321	-0.256	1.746	1.143	-0.836	-0.020	2.224	1.286	1.865	0.120
Peru	Bear Market	-0.080	0.077	0.108	7.005	0.131	0.947	4.443	2.965	5.970	1.636	1.134	0.736
	Bull Market	9.195	0.625	0.364	3.207	3.211	1.402	1.359	0.890	0.321	0.833	1.825	5.476
Venezuela	Bear Market	4.042	3.355	2.868	6.794	0.963	1.623	2.384	3.317	7.418	3.336	1.365	-0.074
	Bull Market	9.521	7.020	0.813	1.342	1.237	-0.327	2.317	0.524	0.083	1.542	1.491	0.937

IV. Conclusion

This study examines month of the year effects for the higher-moments of the returns distribution for Latin American indexes and the effects of diversification on those risk factors. The results provide evidence that kurtosis plays a stronger role in violations of normality than skewness does, and that while violations of normality occur most strongly across the region in the middle of the year (June through September) significant differences in the degree of skewness and kurtosis across the months of the year are sporadic. The study also examines the effects of diversifying across the region for each month of the year and provides evidence that while standard deviation and kurtosis characteristics improve with diversification, skewness of the portfolio becomes more negative, and that this may be due to the fact that the indexes react more quickly to bad news events than to good news events. Moreover, the increased exposure to skewness risk that results from regional diversification is mitigated by decreases in variance and kurtosis risk. Finally, diversification across the region has no meaningful effect on co-variance, co-skewness, or co-kurtosis risk in any month of the year. The exception is the month of September where regional diversification results in a decrease in systematic risk exposure to the third moment of the distribution and provides some mitigation for the exposure to the negative skewness observed in that month.

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Article Review

The Economics of HIV/AIDS in Low Income Countries The Case for Prevention, David Canning, The Journal of Economic Perspectives, Summer 2006.

David Canning is Professor of Economics and International Health, Harvard School of Public Health. He notes there are two approaches to reducing the burden of sickness and death associated with HIV/AIDS, treatment and prevention. Prevention measures limit the transmission of AIDS, and include mass media campaigns, peer education of prostitutes, prevention of mother to child transmission, diagnosis and treatment. The needs for prevention and treatment in low income countries outstrip the resources available. The standard policy prescription is that in order to maximize health with a limited budget, funds should first be allocated to more cost effective interventions and only then to interventions with lesser cost effectiveness.

Among international health agencies, the consensus view is for a rapid expansion of treatment programs in developing countries. The World Health Organization has made the treatment of HIV/AIDS their priority. The United States in its President's Emergency Program for AIDS Relief, a five year $1 billion plan, will focus 80 percent of its resources on treatment and care, with 20 percent going to prevention efforts. Canning argues that the above approaches are a mistake.

There were about 40 million people in the world in 2005 infected with HIV, about 26 million of them in Africa. HIV/AIDS has large potential impacts on the economy and society. Health can be taken to represent a form of human capital, and intervention can boost school attendance and worker productivity.

A number of prevention measures can dramatically lower the transmission of HIV/AIDS. One set of measures focuses on nonsexual pathways, from mothers to children, and through blood transfusions. A second set of prevention measures focuses on changes in sexual activity, promoting abstinence. A third set targets transmission rates for other sexually transmitted diseases. Voluntary counseling and testing is widely available and has a role in prevention strategies.

Overall the evidence for the cost effectiveness of prevention is clear cut for the nonsexual pathways. The effects of intervention that change behavior, or change transmission rates, need to be calculated in an epidemiological model that calculates how incidence and prevalence rates respond to the change. The case for prevention is based to a large extent, on the externality that each infection leads to added transmission to others so that the public benefits of prevention efforts outweigh the private benefits.

Countries with very low incomes face resource limitations. In these countries the issue is to what extent the international community should fund medical therapy. Private charity can of course choose to focus attention on any particular group to which it feels a strong connection. However, if the goal is to maximize the health benefits produced, developing country governments should focus their health spending first on the prevention of HIV transmission, before moving on to treatment.

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