Mean Reversion Between Different Classes of Shares in Dual-Class Firms: Evidence and Implications *

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Abstract

This paper analyzes the time variation of relative prices of stocks with differential voting rights in the world. In the United States, I find that out of 101 firms with at least two classes of stocks traded publicly, 81 (80%) are fractionally cointegrated. The mean (median) estimated long-memory parameter of the equilibrium is 0.61 (0.65) and is negatively correlated with the quality of the firm's corporate governance. The fractional cointegration among dual-class shares implies mean-reverting residual process if long-memory parameter is less than 1. It also implies predictability of future returns using the past prices. A simple long-short strategy based on the cointegration regression residuals yields substantial abnormal returns, even after controlling for market, size, book-to-market, momentum, liquidity, and investor sentiment. Extending the analysis to the rest of the world, I find that out of 788 firms with dual-class shares traded publicly, 470 (60%) are fractionally cointegrated. The long-memory parameter for a country is negatively correlated with quality of its investor protection.

JEL Classification. F30, G15, G34
Keywords. Dual-class shares, Voting premium, Fractional cointegration

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1 Introduction

The principle of "one-share one-vote" (i.e. proportional alignment of cash-flow rights with voting rights) has generated much interest in the literature and the public debate.[1] However, the literature is still inconclusive whether one-share one-vote is beneficial or detrimental to the value of a company. Adams and Ferreira (2008) and Burkart and Lee (2008) provide an excellent discussion of the issue from empirical and theoretical perspectives, respectively.

One of the important phenomena studied extensively in the one-share one-vote debate is the firms with multiple classes of common stock with differential voting rights (hereafter referred to as the "dual-class" firms[2]). In the US, six percent of publicly traded companies are dual-class. A typical dual-class company has an "inferior" class of stock with one vote per share and a "superior" class of stock with ten votes per share (Gompers, Ishii, and Metrick (2009)). In other countries, the voting ratio[3] might be different from 1:10. For instance, in Brazil, Germany, and South Korea the voting ratio is 0:1, whereas in South Africa it ranges from 1:100 to 1:500 (Doidge (2004)).

Dual-class firms provide an excellent tool to study the effects of (dis)proportional ownership since their shares with differential voting rights can be used to easily disentangle the cash-flow and voting rights attached. For instance, Gompers, Ishii, and Metrick (2009) use the dual-class shares to unbundle the incentive and entrenchment effects in the relationship of insider ownership and firm value. They find that firm value is positively (negatively) associated with insiders' cash-flow (voting) rights and negatively associated with the wedge between the cash-flow and voting rights. Another example

[1] In 2006, the European Commission initiated an effort to understand the effects of (dis)proportional ownership and to use this information in its future policies to regulate the capital markets in the European Union (EU). The study was led by a consortium Institutional Shareholder Services (ISS), the European Corporate Governance Institute (ECGI) and Shearman & Sterling LLP. After reviewing the arguments advanced, the Commission has decided in December 2007 that there is no need for action at EU level on this issue since the studies do not provide conclusive evidence on the effect of (dis)proportional ownership (see http://ec.europa.eu/internal_market/company/shareholders/indexb_en.htm for the details of the study and the reports produced).

[2] Although rare, some firms have more than two classes of shares. In the paper, for convenience, they will still be referred as "dual-class" firms. For these firms, the two classes of shares where the difference between voting rights are the largest will be used in the analysis throughout the paper.

[3] Voting Ratio (r) = Voting rights per inferior share / Voting rights per superior share.

where dual-class firms are very useful in analysis is the studies about the value of control in corporations. For instance, by studying the premium paid for superior shares (i.e. voting premium[4]) across 18 countries, Nenova (2003) finds that average voting premium in a country is negatively correlated with its quality of investor protection.

In studies with dual-class shares, the voting premium is usually measured by averaging the daily premium in a given time period (e.g., a year) and is implicitly assumed to be stationary around a constant without any proper justification of the assumption. In other words, previous research focuses on the voting premium in a "static" way and the dynamics of the premium during time has not been analyzed thoroughly. However this static measure might bias the inference about the value of control depending on the time series properties of the data. Using a new method based on option prices to measure the value of voting rights, Karakaş (2009) provides evidence for a considerable time variation in the value of voting rights. He finds that the value of voting rights is in general negligible, but can be quite high when the control is contested. A careful investigation on the time variation of dual-class share prices would allow for a better estimation of the voting premium and understanding of the determinants of the voting premium.

This paper studies the time variation of relative prices of dual-class shares using two different sets of data. Employing the comprehensive US dual-class firms data constructed by Gompers, Ishii, and Metrick (2009), I find that 81 out of 101 dual-class firms, where at least two classes of stocks traded publicly, are fractionally cointegrated. Two random walks are called fractionally cointegrated, if there exists a linear combination of the two series which is not a random walk but a long-memory process. Using a world-wide data set constructed by Doidge (2004), I find that 60% of the dual-class shares are fractionally cointegrated in the rest of the world. The average long-memory parameter estimates for the equilibrium error process is 0.61 for US and 0.64 for the rest of the world, which implies that the equilibrium errors are not stationary but mean-reverting.

According to Granger (1987), if two stocks are cointegrated then future returns of at least one of them is predictable by the past prices of both series. Following this

[4]Formal definition of voting premium is provided in Section 2.1.

argument, I propose and analyze a simple long-short trading strategy based on the cointegration regression residuals of the US sample. I find that this strategy leads to substantial abnormal returns (around 14% annual excess return), even after controlling for widely used risk factors, such as market, size, book-to-market, momentum, liquidity, and investor sentiment. Market, size and value factors are not significant in explaining the excess return, most probably due to the dual-class structure of the stock pairs (i.e. due to going short and long fundamentally the same stock). Momentum factor is significant and negative. This is consistent with the trading strategy since it bets on the mean-reversion.

The fractional cointegration relation among the dual-class shares has also governance implications. For the US sample, at the company level analysis, I find that the better the firm's quality of governance (measured by the governance index proposed by Gompers, Ishii, and Metrick (2003)) the smaller its long-memory parameter. Note that smaller long-memory parameter implies faster convergence for the mean-reverting residual process. Therefore the result suggests that the worse the governance of a firm, the longer it takes for the prices of different classes of shares to converge to their long-run equilibrium. In a similar fashion, for the world sample, at the country level analysis, I find that the better the country's investor protection (measured by revised anti-director and anti-self dealing indices proposed by Djankov, Porta, de Silanes, and Shleifer (2008)) the smaller the average long-memory parameter of the dual-class firms in the country.

This paper extends the study by Dittmann (2001). He analyzes the voting and non-voting shares of ten German companies and finds that seven pairs of price series are fractionally cointegrated with the long-memory parameter of the equilibrium errors lying between 0.5 and 0.8. Despite the small size of his sample, my results are consistent with his findings. The main differences between our papers are as follows: First, he only focuses on a sample of ten dual-class companies in Germany whereas I cover all dual-class firms (101 in US sample and 788 in the rest of the world) in 21 countries. Second, his trading rule has perfect foresight whereas mine does not look ahead. Third, in his trading results he only reports returns without a control for risk whereas I control for widely used risk factors. Finally, he does not investigate the relation of long-memory parameter to the governance of a company or to investor protection of a country, whereas

I do.

The paper proceeds as follows. Section 2 provides some background about voting premium and fractional cointegration, and briefly reviews the related literature. Section 3 describes the data. Section 4 discusses the methodology and results. In Section 5, a simple trading strategy is introduced and analyzed. Section 6 presents additional results about corporate governance and investor protection. Finally, Section 7 concludes.

2 Background

2.1 Voting Premium

As mentioned in the introduction, dual-class firms are very useful in the studies about private benefits of control in a corporation. By very definition, private benefits of control are very difficult to measure. In order to be able to infer the value of control there are two main methods proposed in the literature.

The first method focuses on privately negotiated transfers of controlling blocks in publicly traded companies and compares the price per share paid by the acquiring party to the price per share prevailing on the market (see, e.g., Barclay and Holderness (1989) and Dyck and Zingales (2004)).

The second method simply studies the market value of different classes of shares. (see, e.g., Lease, McConnell, and Mikkelson (1983), Lease, McConnell, and Mikkelson (1984), Nenova (2003), Zingales (1994), and Zingales (1995)).[5]

Compared to the first method, it is easier to collect the data in the second method. Moreover it allows more freedom to do analysis about the value of control at any time period since it does not necessarily require a transfer of controlling blocks. Since the same company's shares are used in the analysis, in most cases, it also allows to do natural experiments and/or to naturally control the unobserved company-specific effects. However studies with dual-class firms also have some drawbacks such as certain classes of shares (usually the superior ones) might be non-publicly traded shares. Even though

[5]Nenova (2003) proposes a crossover method where she infers the control block (assumed to be 50% of the voting power) premium from the difference between the market value of dual-class shares.

both classes of voting shares are traded, one class might be less liquid and/or the price differential among different classes of shares might reflect the valuation of a minority investor. More importantly, the studies might be subject to endogeneity issues such as private benefits of control are higher in dual-class firms compared to the firms following one-share one-vote principle (see, e.g., DeAngelo and DeAngelo (1985) and Smart and Zutter (2003)). In a recent article, Karakaş (2009) proposes a new technique of inferring the value of control by unbundling the value of cash-flow and voting rights in shares using derivative products (e.g. option contracts). This method might help solve the endogeneity problems of dual-class studies since it facilitates using both the single- and dual-class companies to determine the value of control.

The relation between the prices of inferior and superior shares in dual-class companies are usually expressed with the voting premium, which is basically defined as follows[6]:

$$VP = (P_S - P_I)/P_I \tag{1}$$

where VP is voting premium, P_S is the price of the superior voting share and P_I is the price of the inferior voting share.

One might be tempted to conclude that the voting premium is always positive since the the right to vote can be seen as an "option". However, this is not the case in reality. Lease, McConnell, and Mikkelson (1983) find that for four out of 30 US companies the superior voting shares trade at a significant discount relative to the inferior voting rights. This suggests that there are both benefits and costs of corporate control and the sign of voting premium depends on which side predominates. There might several reasons for negative voting premia. For instance, Neumann (2003) suggests that liquidity risk explains the negative premia observed in Denmark during the period 1992 to 1999 and Ødegaard (2007) argues that restrictions on foreign ownership on voting shares cause the negative premia in Norwegian market between 1988 and 1994.

[6]Note that this formula does not take the voting ratio of the dual-class firm which might create bias in cross-section analysis with firms having different voting ratios. A more consistent formula (offered by Zingales (1995)) taking this issue into account is:

$$VP = (P_S - P_I)/(P_I - rP_I)$$

where r is the voting ratio. As easily observed, the formula collapses to the one in Equation 1 if r is zero.

The time-series variation of the voting premium has not been studied extensively. The literature until now has mostly focused on the cross-section variation in the voting premium. Zingales (1995) shows that the voting premium is determined by the expected additional payments vote holders will receive if there is control contest in US firms. Using 661 dual-class firms in 18 countries in 1997, Nenova (2003) demonstrates that value of control block varies widely across countries (e.g. average ratio of value of control-block votes to firm value is 29% in Italy whereas it is 2% in the US) and control-block votes are significantly less value in a stricter legal environment (e.g. average value of control-block votes is 4.5% in Common Law countries, whereas it is 25.4% in French Law origin countries.). Using 137 out of 745 non-US dual-class firms cross-listed in the US over the period from 1994 to 2001, Doidge (2004) finds that voting premia are 43% lower for non-US firms that choose to cross-list on US exchanges compared to the non-US firms that do not cross-list. This supports the bonding hypothesis that controlling shareholders can use a US listing to bond themselves to assure minority shareholders that they are less likely to be exploited.

2.2 Fractional Cointegration

Let S_t and I_t denote the prices of superior and inferior voting shares at time t, respectively. Assume that log prices of superior and inferior voting (denote s_t and i_t, respectively) shares follow a random walk with uncorrelated increments[7]:

$$s_t = \ln S_t \quad and \quad i_t = \ln I_t \tag{2}$$

$$s_t = r^s + s_{t-1} + \varepsilon_t^s \quad and \quad i_t = r^i + i_{t-1} + \varepsilon_t^i \tag{3}$$

where ε_t^s and ε_t^i are uncorrelated error processes with zero mean.

s_t and i_t are *fractionally cointegrated*, if there exist constants κ, β and $\gamma \in R$, such that:

$$i_t = \kappa + \beta s_t + \gamma t + u_t \tag{4}$$

[7]This assumption will be checked in the following empirical tests and the stocks that violate this assumption will be out of the sample.

where u_t is a long-memory or $I(d)$ process with long-memory parameter $d \in (0, 1)$. u_t is stationary if $d < 0.5$. See Baillie (1996) for a survey and review of econometric literature on long-memory process.

If long-memory parameter $d = 0$, s_t and i_t are classically cointegrated. If $d = 1$, then u_t is integrated and therefore s_t and i_t are not cointegrated. Fractional cointegration stays between "no cointegration" and "cointegration" cases.

Note that an $I(d)$ process with $d < 1$ is a mean-reverting process (Cheung and Lai (1993)). This suggests that any shock to the system dies out, albeit slowly, and therefore the Equation 4 defines an equilibrium if u_t is an $I(d)$ process with $d < 1$.

In the literature, fractional cointegration has been used as a tool in many different studies (e.g., exchange rates (Baillie and Bollerslev (1994)), purchasing power parity (Cheung and Lai (1993)), and stock returns (Lo (1991))). Dual-class shares are very suitable for the fractional cointegration analysis since for each pair of different classes of shares, the underlying company is the same. This is an advantage compared to other studies trying to find common property among different assets which are not necessarily tied to each other fundamentally.

3 Data

3.1 US Sample

I use the data compiled by Gompers, Ishii, and Metrick (2009). In their paper, they establish a comprehensive list of dual-class shares in the United States between 1995 and 2002. The average number of dual-class firms is 450 per year. Since I need the price data for both superior and inferior voting shares of a firm, I use a subset of their data where at least both classes of shares are publicly traded. This drops the average number of dual-class firms to 70 per year. I get the price and the volume data from the CRSP. After recognizing the dual-class shares between 1995 and 2002, I have enlarged the sample data by going back/forth in time as much as possible (from 1980 to 2006) since fractional cointegration needs long time series to be detected. The analysis below are based on this enlarged sample. However, the results are qualitatively the same for the original sample of Gompers, Ishii, and Metrick (2009).

Following Dittmann (2001), I constructed weekly price data of superior and inferior shares since missing values due to holidays, weekends and/or no trading might distort the periodogram. The construction of the the weekly observations is as follows:

- Price series are adjusted for dividends, stock splits etc. dividing the raw price data by the "Cumulative Factor to Adjust Price - CFACPR" variable in the CRSP.
- Observations with zero trading volume ("Share Volume - VOL" variable in the CRSP is equal to zero) are discarded.[8]
- For each individual series, weekly observations are generated by selecting the mid-week observations to avoid non-trading, bid-ask spread, asynchronous prices biases and the beginning/end of the week anomalies inherent in daily data. For mid-week observation, Wednesday observation is selected. If it is missing then Tuesday's price (then Thursday's, Monday's, Friday's) observation is used instead.
- Superior and inferior stock series are merged. After all I have a total sample of 101 companies.

3.2 World-Wide Sample

I use the data provided by Doidge (2004) which covers 20 countries. Using a similar procedure for the US firms described above, I construct weekly prices. Slight differences in the procedure is as follows. First, instead of CRSP, I use Datastream to get the daily price and volume data of the shares. Second, I merge inferior and superior prices before constructing the weekly prices. This ensures to have synchronous prices. Finally, since the volume data is not of good quality, it is not used in the analysis.

The steps in construction of the weekly observations and the remaining number of companies in the dataset after each step are as follows:

- Initial sample: 798 companies from 20 countries (exluding US).

[8]When the volume data is missing the observations are not discarded since it is not clear whether the volume is zero or not. Occasionally the price changes even though the volume observation is missing.

- Clean problematic data: 788.
- Drop companies whose prices are not random walk process individually: 624.
- Drop companies whose prices are not cointegrated: 470.
- Drop companies whose long-memory parameter is not between 0 and 1, and whose parameter is not significant at 5% level: 411.[9]
- Add US data, constructed at previous section, after dropping the non-significant long-memory parameters: 480 companies from 21 countries.

4 Methodology & Results

4.1 US Sample

I mostly follow the methodology discussed in Dittmann (2001). First, I test each individual series for being a random walk using Phillips-Perron test with lag truncation parameters 0, 5, 10, 15, 20 and $\text{int}\{4(T/100)^{(2/9)}\}$ where T is number of observations. I exclude the companies from further analysis if the random walk hypothesis is rejected at the 5% level for any class of their shares at any of the tests. This reduces my sample from 101 to 88 companies.

Second, I test for cointegration relationship between the superior and inferior voting shares in the dual-class companies. The null hypothesis is that s_t and i_t are not cointegrated, i.e. u_t in Equation 4 is $I(1)$ for any κ, β or γ. In a first step, the following regression is estimated by OLS:

$$i_t = \hat{\kappa} + \hat{\beta} s_t + \hat{\gamma} t + \hat{u}_t \quad (5)$$

In the second step, the Phillips-Perron unit root test is applied to the regression residuals $\hat{u}_t$ with lag truncation parameters 0, 5, 10, 15, 20 and $\text{int}\{4(T/100)^{(2/9)}\}$ where T is number of observations.[10] For 81 companies (out of 88), the test rejects the hypothesis of no cointegration at the 5% level at any lag specification. Table 1 and Table

[9]Results are robust to skipping this step.

[10]See Hamilton (1994) for details.

2 display the overview of the Phillips-Perron test statistics for the full-sample (including not cointegrated companies) and the reduced-sample (excluding not cointegrated companies). Results clearly demonstrate that 81 of the initial 101 pairs of price series are cointegrated.

[Insert Table 1 & 2 about here]

Table 3 and Table 4 display the overview of the OLS estimates for cointegration regression (Equation 5) for the full-sample (including not cointegrated companies) and the reduced-sample (excluding not cointegrated companies).

[Insert Table 3 & 4 about here]

Note that average $\hat{\beta} \approx 1$ and average $\hat{\gamma} \approx 0$. Therefore the voting premium is

$$VP = \frac{P_S - P_I}{P_I} = \exp((1-\beta)v_t - \kappa - \gamma t) - 1 \approx \exp(-\kappa) - 1 \approx -\kappa \tag{6}$$

Therefore $-\hat{\kappa} = 11.4\%$ can be regarded as the average voting premium for the sample. This magnitude is similar to the average voting premium reported in Zingales (1995) for the United States, which is 10.47%.[11] However, the Fama-Macbeth (FM) type t-statistic implies that $-\hat{\kappa}$ is not significantly different than zero whereas it is significantly positive in Zingales (1995).[12] This is because in my sample the variation is very high with stdev($\hat{\kappa}$) around 77% whereas the variation is relatively smaller in Zingales (1995) with standard deviation of the voting premium around 24%. This is not very surprising since Zingales (1995) calculates the voting premium by taking the annual average of daily voting premium over the seven years (from 1984 to 1990) for each of 94 companies in his sample, which would clearly reduce the standard deviation.

Table 5 and Table 6 display the overview of three Geweke and Porter-Hudak (1983) estimates of the long-memory parameter, $\hat{d}$, of cointegration regression residuals $\hat{u}_t$ for

[11] Note that in Zingales (1995) voting premium is defined as $(P_S - P_I)/(P_I - rP_S)$ where r is the relative number of votes of an inferior voting share versus a superior voting. In my sample r is usually 0.1. Incorporating that back-of-the-envelope calculation yields $-\hat{\kappa}' \approx 12.8\%$.

[12] In untabulated results using the original sample of Gompers, Ishii, and Metrick (2009), $-\hat{\kappa} \approx 3\%$ and again not significantly different than zero according to Fama-Macbeth type t-statistic.

the full-sample (including not cointegrated companies) and the reduced-sample (excluding not cointegrated companies).

[Insert Table 5 & 6 about here]

The mean estimated long-memory parameter of the equilibrium is around 0.61. About 68% of the companies have $\hat{d}$ greater than 0.5, which implies that the estimated residual process $\hat{u}_t$ is not stationary.

4.2 World-Wide Sample

Following the same methodology for the US sample, I find that out of 788 dual-class companies (excluding US) 164 have at least one class of share which is not following random walk process at 5% level and are excluded from further analysis. For 470 companies out of remaining 624, the Phillips-Perron test rejects the hypothesis of no cointegration at 5% level.

Table 7 displays the overview of the OLS estimates for cointegration regression (Equation 5) for the world-wide sample including the US sample. As observed in the table, there is quite a lot of variation in the estimates, especially in $\hat{\kappa}$ and $\hat{\beta}$, across countries. For the countries with average $\hat{\beta} \approx 1$ and average $\hat{\gamma} \approx 0$ (e.g., Canada and South Africa), the average voting premiums, which can be proxied by $-\hat{\kappa}$ as shown in Section 4.1, are quite close to the ones reported in Table 2 of Doidge (2004). For other countries where average $\hat{\beta}$ is not close to one, the concept of voting premium is not very clear. As shown by Dittmann (2001), β is a parameter showing the relative riskiness of the different classes. If β is smaller than one, voting shares are more risky than the non-voting shares in the long run and vice versa if β is greater than one. In most of the countries (18 out of 21), the $\hat{\beta}$s are close to or less than one and therefore the voting shares are more risky compared to non-voting ones. In remaining three countries, namely Australia, Austria and Mexico, it is the other way around. Note, however, that Australia and Austria have very small samples that might drive this result.

5 Trading Strategy

In previous section, I show that substantial fraction of dual-class shares (80% in the US and 60% in the rest of the world), whose at least two classes of shares publicly traded, are fractionally cointegrated. The fractional cointegration among dual-class shares implies predictability of future returns using the past prices. In order to check whether this anomaly is economically significant, I evaluate a simple trading strategy proposed by Dittmann (2001) using the US sample. The reason I focus on the US sample is that this gives the advantage of easily accessing the widely-used risk factors namely market, size, book-to-market, momentum, liquidity, and investor sentiment.

Consider the following trading strategy:

- If $\hat{u}_t$ drops below -0.05, sell (buy) voting (non-voting) shares.
- If $\hat{u}_t$ rises above 0.05, buy (sell) voting (non-voting) shares.

Note from Equation 5 that if $\hat{u}_t$ drops below -0.05, non-voting share is relatively cheaper. The trading strategy is therefore to long the "cheap" non-voting share and to short the "expensive" voting share. Similarly, trading strategy is to long the "cheap" voting share and to short the "expensive" non-voting share once $\hat{u}_t$ rises above 0.05. This is a strategy which bets on the mean reversion. However, remember that long-memory parameter is almost always less than 1 in the sample (see Table 5) and therefore the process is mean-reverting (see Section 2.2). Note also that under perfect foresight every time $\hat{u}_t$ crosses -0.05, +0.05 lines, the trading strategy yields a 10% return.

In order to have a more realistic trading rule, for each company, starting from the 53rd week (after 1 year of data) I use the past data to calculate $\hat{u}_t$ every week. Then according to the calculated $\hat{u}_t$, I apply the trading strategy described above. This is an implementable trading strategy in the sense that no foresight is used.[13]

[13] One caveat is that in the construction of the weekly prices, following Dittmann (2001) I get the mid-week price of each class of share separately trying to be as close to Wednesday as possible. This sometimes creates non-synchronous prices for a given week. For instance, if the Wednesday price data at a certain week is available for a non-voting share but is not available for a voting share (assuming that the price data exists for voting share in another day at the week) then that week's weekly prices will not be synchronous. This in part might create a slight foresight. I am currently working on this issue to check whether it changes the results qualitatively.

5.1 Fama-French + Carhart Four Factors

Table 8 displays the overview of Fama-French + Carhart four factor regression results. The regressions are done with monthly returns. According to Fama-Macbeth type t-statistics there seems to be a substantial *Alpha* (around 1.2%). This corresponds to about 14% yearly excess return. The other factors do not seem to be significant.

Table 9 again displays the overview of Fama-French + Carhart four factor regression results but for only companies whose number of observations are greater than 50. Similar to the results observed at Table 8, there appears to be a substantial *Alpha*. Note that in this sub-sample, the momentum factor (UMD) is also significant and negative in sign. This is consistent with the trading strategy since it bets on the mean-reversion. Therefore when there is momentum it affects the returns negatively. However, even though it is significant the *Alpha* is still around 1.2%, which is economically significant.

Table 10 displays the Fama-French + Carhart four factor regression results for the equal-weighted calendar-time portfolios formed monthly.[14] Similar to results in Table 8 and 9, there is an excess *Alpha* of 1.1% which is highly significant. Moreover momentum factor negatively affects the returns as seen at Table 9.

Overall, the results indicate that the cointegration relationship between different classes of dual-class shares demonstrated in previous sections can be used to predict the future returns. A simple long-short strategy leads to around 14% annual excess return.

14% annual excess return is a plausible figure. For instance, using a simple "pairs trading" strategy, Gatev, Goetzmann, and Rouwenhorst (2006) document an average annualized excess returns of up to 11% for self-financing portfolios formed between 1962 and 2002. Pairs trading strategy is basically finding two stocks whose price move together historically and betting on the convergence of the prices once the spread between them widens by going long (short) the loser (winner). From pairs trading perspective, one can interpret dual-class shares as "natural" pairs. Another recent example is the study by Kosowski, Naik, and Teo (2007) documenting that top decile hedge funds persistently generate about 5% excess alpha per annum compared to the bottom decile funds using

[14] See, e.g., Mitchell and Stafford (2000) for a discussion of calendar-time portfolios (and also other methods, to measure long-term stock price performance).

net-of-fee returns of live and dead hedge funds reported in the union of TASS, HFR, CISDM, and MSCI databases which represent the largest known data set of hedge funds to date.

5.2 Liquidity

Usually the voting shares are more illiquid. In order to check whether liquidity might explain the excess returns, I run the time series regressions with the Pastor-Stambaugh liquidity measures. I tried "Liquidity Level (non-traded factor)", "Liquidity Innovation (non-traded factor)", "Equal-Weight Traded Factor", and "Value-Weight Traded Factor", which are downloaded from WRDS. In untabulated results, I find that neither of these factors are significant in any regression. The *Alpha* is always significant and around 1.1%. Momentum factor, UMD, is significant with negative sign when sample is restricted for companies with number of observations greater than 50.

As for another factor, I use the liquidity factor provided by Viral Acharya. The limitation of this data is it is until the end of year 1999. Running the time series regressions with this liquidity factor added, in unreported results I find that it is not significant according to Fama-Macbeth type t-statistics. In regressions with calendar-time portfolios, however, the liquidity factor enters significantly with a positive sign (see Table 11). This however does not affect the significance and the magnitude of *Alpha*, which is still highly significant and around 1.3%. Note that the HML factor is also significant with a negative sign. However, this seems to be more of a period dependent event rather than the inclusion of the liquidity factor. This is because I run the same regression without liquidity factor and find that HML factor has very similar figures for the significance and the magnitude.

Overall the the market-wide liquidity factors (e.g. Pastor-Stambaugh) do not seem to explain the excess *Alpha*. In fact, in most of the cases the liquidity factors are not even significant. However this still does not rule out the possible role of liquidity in explaining the reported excess returns. The individual illiquidity of the stocks -rather than the market-wide liquidity factors- is very important and requires further analysis.

5.3 Investor Sentiment

Using the monthly data published by Jeffrey Wurgley at his homepage[15], I checked whether investor sentiment might explain the excess *Alpha*. In unreported results, I find that investor sentiment factor is not significant in the regressions and the results are qualitatively the same as previously reported ones. Overall, investor sentiment does not appear to explain the anomaly.

6 Corporate Governance & Investor Protection

6.1 Corporate Governance

How does governance of the company relate to the cointegration relationship between the superior and inferior voting shares?

Using the data from US sample, Figure 1 displays that there is a positive correlation between estimated long-memory parameter ($estd$) and the governance index (g), which is downloaded from WRDS. The governance index is a measure of a firm's quality of corporate governance developed by Gompers, Ishii, and Metrick (2003). The index is the count of corporate-governance provisions that reduces shareholder rights in a company's charter, bylaws, and/or SEC filings. Therefore the higher the index (max. 24), the worse the quality of corporate governance. See Gompers, Ishii, and Metrick (2003) for detailed discussion of the construction of the governance index. Note that since the governance index does not exist for some of the companies, the sample decreases from 81 to 45 companies.

Although the governance index (measured once in 2-3 years) is not static, it is quite stable. It mostly stays same or goes up/down by one. In other words, time variation almost does not exist. Therefore to have a cross-section regression, for each company I take their average governance index and regressed the estimated long-memory parameter on it. Figure 1 shows the fitted values and the 95% confidence interval for the regression. Table 12 displays the results of the regression. Correlation between long-memory

[15] http://pages.stern.nyu.edu/ jwurgler/

parameter and the governance index is 0.356 and significant (p=0.017).

This indicates that as governance index goes up (which means that the quality of corporate governance goes down), the long-memory parameter increases. This is not necessarily a causal relationship. Note that the higher the long-memory parameter, the longer it takes to converge to the mean for the residual process. Therefore the worse the governance is, the longer to converge to the long run equilibrium mean for the residual process. One interpretation of this result might be that as the governance becomes worse, the market becomes less efficient in processing the information and/or the company itself goes through longer periods where the divergence of the prices occur.[16] This result is consistent with the findings of Ali, Chen, and Radhakrishnan (2007) that family firms tend to disclose less information about their corporate governance practices in their proxy statements and family firms with dual-class shares have more agency problems between controlling and non-controlling shareholders compared to the ones without dual-class shares. Villalonga and Amit (2006) also find evidence that family firms with dual-class shares have more severe agency problems than the ones without dual-class shares.

One might expect the governance index to be positively correlated with $\hat{\kappa}$ since I show that $\hat{\kappa}$ is measuring voting premium, which can be interpreted as the lower bound of the private benefits. However in my sample governance index and $\hat{\kappa}$ are not significantly correlated (unreported).

Figure 2 shows the the fitted values and the 95% confidence interval for the regression of *Alpha* on estimated long-memory parameter. Table 13 displays the results of the regression. Correlation between *Alpha* and long-memory parameter is -0.435 and significant (p=0.000).

This indicates that *Alpha* and long-memory parameter are negatively correlated.

[16] Lee, Myers, and Swaminathan (1999) make a similar argument in their paper modeling the time-series relation between price and intrinsic value as a cointegrated system, so that price and value are long-term convergent. They state: "The case for the equality of price and value is based on an assumption of insignificant arbitrage costs. When information and trading costs are trivial, stock prices should be bid and offered to the point where they fully reflect intrinsic values. However, when intrinsic values are difficult to measure and/or when trading costs are significant, the process by which price adjusts to intrinsic value requires time, and price does not always perfectly reflect intrinsic value. In such a world, a more realistic depiction of the relation between price and value is one of continuous convergence rather than static equality".

This result is intuitive since the trading strategy is designed to make 10% (under perfect foresight) every time it crosses ± 0.05. The larger the long-memory parameter the longer it takes to cross due to slow mean reversion and therefore the lower the *Alpha* is.

Based on the findings that there is positive correlation between governance index and long-memory parameter and there is negative correlation between long-memory parameter and *Alpha*, I check whether governance is negatively correlated with *Alpha* but I do not find a significant correlation in my sample (unreported).

6.2 Investor Protection

Using the data from world-wide sample, Figure 3 displays that there is a negative correlation between estimated long-memory parameter (*estd*) and the investor protection ($Revised\ Anti-Director$),[17] which is coded by Djankov, Porta, de Silanes, and Shleifer (2008). Anti-director index is the count of six important shareholder rights (e.g., vote by mail, oppressed minority etc.) in a country. It is a measure of a country's quality of investor protection. The higher the index, the better the quality of investor protection in the country. In a similar spirit, anti-self dealing is another measure of investor protection. It has been developed recently to address the concerns about the anti-director rights index. See Djankov, Porta, de Silanes, and Shleifer (2008) for detailed discussion of the construction of both measures.

The negative correlation between long-memory parameter and the investor protection indicates that for a country if the quality of investor protection is higher then the long-run memory parameter is smaller. Therefore for countries with better investor protection, the convergence of residual process of the dual-class share prices is faster. One might consider this finding similar to the finding about corporate governance at company-level analysis in the previous section but at the country-level.

Following Djankov, Porta, de Silanes, and Shleifer (2008), I classify countries according their origin of law in Table 14.[18] English origin countries have the lowest average long-memory parameter (0.575), whereas French origin countries have the highest aver-

[17] Using anti-self dealing index as a measure of investor protection gives qualitatively the same result.

[18] In Table 14, "Sample fw" stands for the sample with frequency-weighted data, whereas "Sample ew" stands for the sample with equal-weighted data.

age long-memory parameter (0.691). In a similar fashion, Common Law countries have lower long-memory parameter compared to Civil Law countries (0.575 vs. 0.666). Results from comparison of means in Table 15 suggest that these difference are significant. I also included the t-stats for the differences between the revised anti-director index of the countries from Djankov, Porta, de Silanes, and Shleifer (2008) in a separate column for comparison. The negative correlation between the long-memory parameter and the quality of investor protection can again be observed clearly.

7 Conclusion

In this paper, I analyze the fractional cointegration among differential voting shares in the world. In the US sample, I find that around 80% of the dual-class shares, whose at least two classes of stocks are publicly trade, are fractionally cointegrated with a mean estimated long-memory parameter of the equilibrium of 0.61. This implies predictability of the future returns. I analyzed a simple long-short trading strategy based on the cointegration regression residuals and find that it yields considerable abnormal returns even after controlling for as market, size, book-to-market, momentum, liquidity, and investor sentiment. Extending the analysis to the rest of the world, I find that around 60% of the dual-class shares are fractionally cointegrated. Long-memory parameter of the process is negatively correlated with the quality of corporate governance at company-level analysis in US sample and is negatively correlated with the quality of investor protection at country-level analysis in the world-wide sample.

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Figure 1: Long-Memory Parameter vs. Governance Index

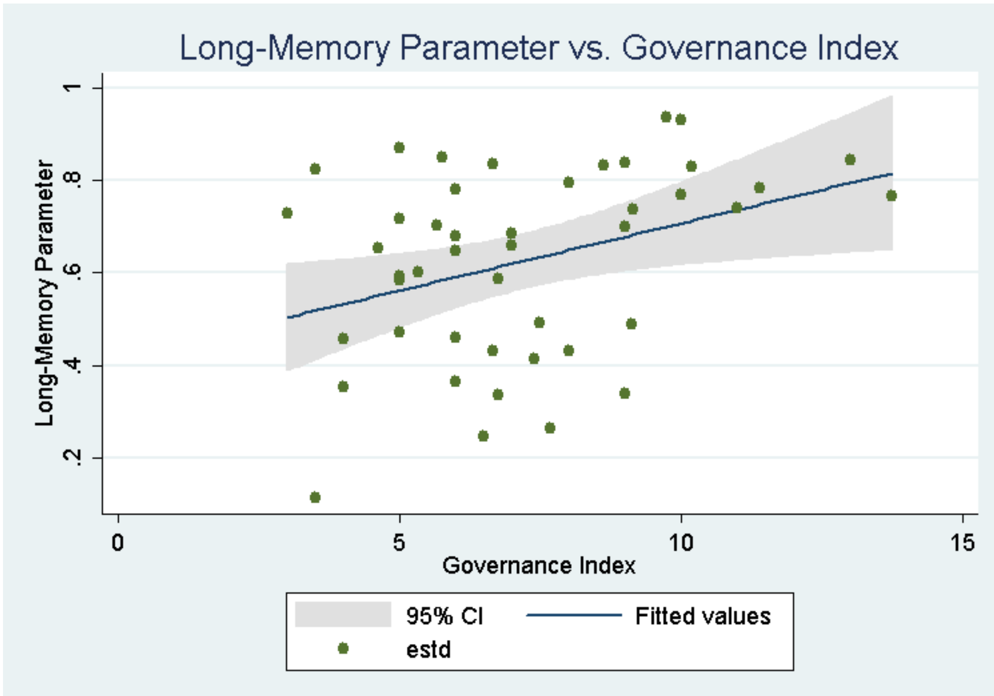

Figure 2: Alpha vs. Long-Memory Parameter

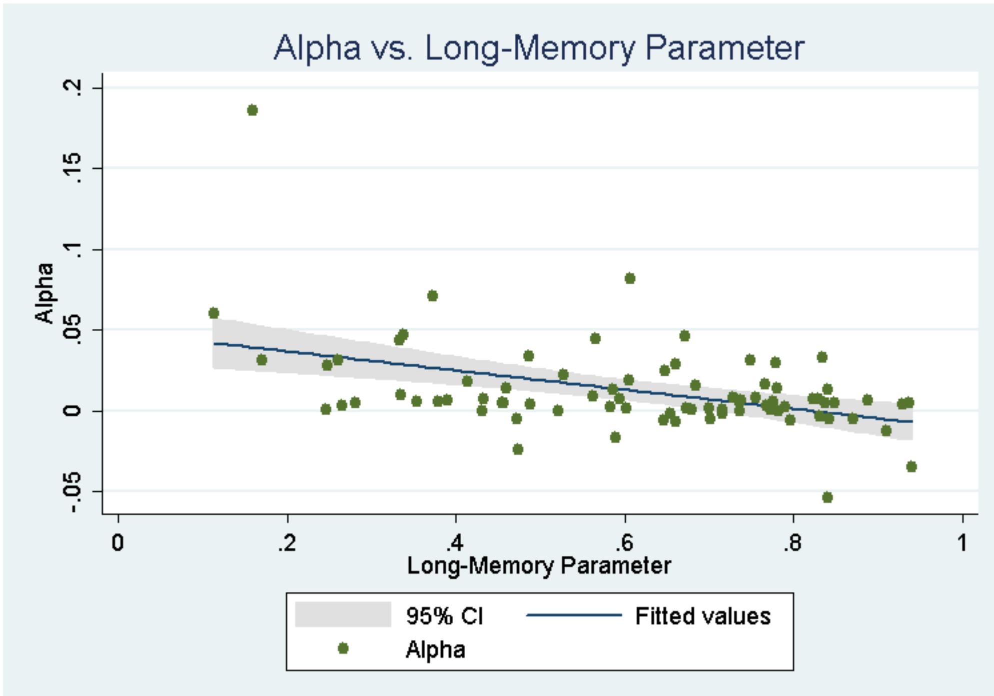

Figure 3: Long-Memory Parameter vs. Investor Protection

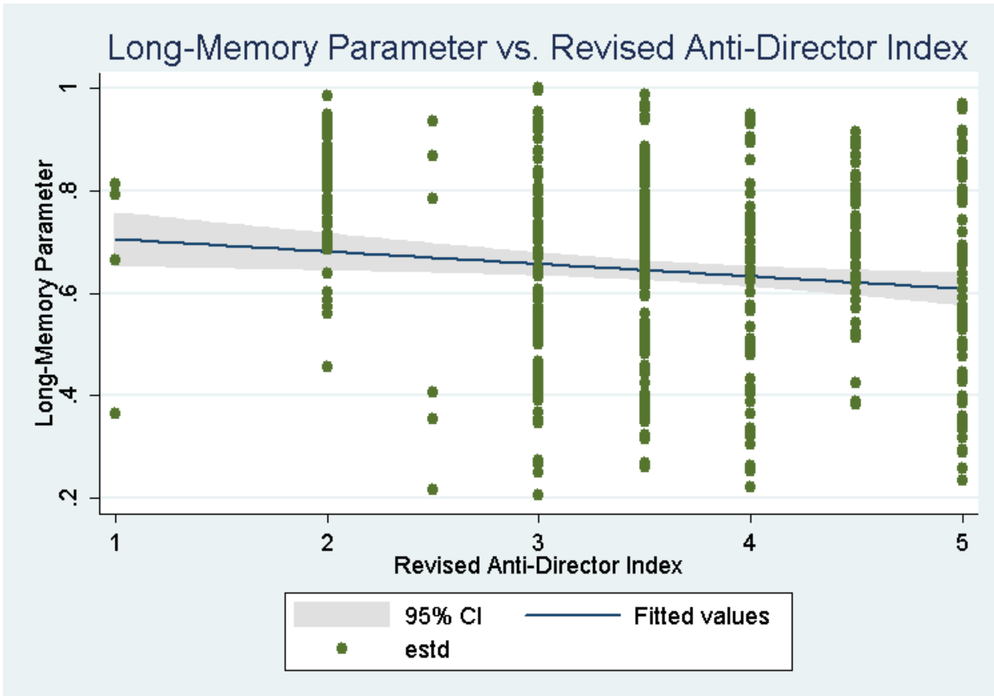

Table 1: Phillips-Perron test statistics of the full-sample (88 companies) including companies with no cointegration. The lag truncation is $\text{int}\{4(T/100)^{(2/9)}\}$ where T is number of observations.

Var	Obs	$Mean$	Std	Min	Max
p	88	0.016	0.064	0.000	0.460
Z_t	88	-8.380	5.086	-24.947	-1.644
Z_ρ	88	-147.373	173.741	-943.139	-8.269

Table 2: Phillips-Perron test statistics of the reduced-sample (81 companies) not including companies with no cointegration. The lag truncation is $\text{int}\{4(T/100)^{(2/9)}\}$ where T is number of observations.

Var	Obs	$Mean$	Std	Min	Max
p	81	0.001	0.006	0.000	0.045
Z_t	81	-8.899	4.968	-24.947	-2.899
Z_ρ	81	-158.974	176.389	-943.139	-14.210

Table 3: OLS estimates for cointegration regression (Equation 5) of the full-sample (88 companies) including companies with no cointegration.

Var	Obs	$Mean$	Std	Min	Max	FM *type t-stat*
$\hat{\kappa}$	88	-0.097	0.795	-3.825	2.479	-1.143
$\hat{\beta}$	88	0.958	0.183	-0.006	1.313	49.020
$\hat{\gamma}$	88	0.000	0.001	-0.005	0.005	1.818

Table 4: OLS estimates for cointegration regression (Equation 5) the reduced-sample (81 companies) not including companies with no cointegration.

Var	Obs	$Mean$	Std	Min	Max	FM *type t-stat*
$\hat{\kappa}$	81	-0.114	0.772	-3.825	2.002	-1.335
$\hat{\beta}$	81	0.967	0.152	0.345	1.313	57.196
$\hat{\gamma}$	81	0.000	0.001	-0.002	0.005	2.506

Table 5: Estimates of the long-memory parameter of the cointegration regression residuals for the full-sample (88 companies) including companies with no cointegration.

	Obs	*Mean*	*Std*	*Min*	*Max*	*Non-Stationary*
$power = 0.50$	88	0.608	0.278	0.088	1.617	59
$power = 0.55$	88	0.622	0.216	0.113	0.996	59
$power = 0.60$	88	0.625	0.228	-0.001	1.198	61

Table 6: Estimates of the long-memory parameter of the cointegration regression residuals for the reduced-sample (81 companies) not including companies with no cointegration.

	Obs	*Mean*	*Std*	*Min*	*Max*	*Non-Stationary*
$power = 0.50$	81	0.600	0.269	0.088	1.617	54
$power = 0.55$	81	0.608	0.208	0.113	0.940	54
$power = 0.60$	81	0.612	0.221	-0.001	1.198	55

Table 7: Estimates of the Long-Memory Parameter of the Cointegration Regression Residuals for the World-Wide Sample.

Country	$\hat{\kappa}$		$\hat{\beta}$		$\hat{\gamma}$		
	Mean	*Std*	*Mean*	*Std*	*Mean*	*Std*	*Freq.*
Australia	-0.373	0.000	1.110	0.000	0.000	0.000	1
Austria	-0.164	2.315	1.109	0.349	0.000	0.001	3
Brazil	0.076	1.323	0.783	0.316	0.001	0.002	54
Canada	-0.124	0.340	0.995	0.111	0.000	0.002	39
Chile	0.094	0.777	0.950	0.186	0.000	0.001	4
Colombia	1.156	0.008	0.820	0.108	0.000	0.000	2
Denmark	0.253	0.786	0.925	0.139	0.000	0.001	21
Finland	-0.485	0.715	0.904	0.195	0.001	0.001	14
France	-0.133	1.295	0.916	0.214	0.001	0.000	6
Germany	0.069	1.097	0.943	0.203	0.000	0.000	51
Italy	-0.466	0.945	0.926	0.226	0.001	0.001	49
Mexico	0.140	0.327	1.128	0.613	-0.001	0.003	11
Norway	-0.216	0.796	1.007	0.067	0.000	0.000	8
Portugal	0.717	0.738	0.776	0.143	0.002	0.001	3
South Africa	-0.054	0.087	0.980	0.046	0.000	0.000	11
South Korea	1.476	4.580	0.861	0.454	-0.001	0.002	57
Sweden	0.323	0.927	0.926	0.165	0.000	0.001	48
Switzerland	0.139	2.128	0.875	0.233	0.000	0.001	10
UK	0.467	0.996	0.904	0.205	0.000	0.001	15
US	-0.124	0.639	0.979	0.138	0.000	0.001	69
Venezuela	0.199	0.258	0.941	0.092	0.000	0.001	4
Total	0.169	1.876	0.922	0.264	0.000	0.001	480

Table 8: Fama-French + Carhart Four-Factor Regressions.

Var	*Obs*	*Mean*	*Std*	*Min*	*Max*	*FM type t-stat*
Alpha	85	0.012	0.030	-0.054	0.186	3.674
MKTRF	85	-0.069	0.740	-3.593	3.106	-0.864
SMB	85	0.145	0.842	-1.484	4.693	1.583
HML	85	0.292	1.948	-1.954	12.164	1.383
UMD	85	-0.142	0.977	-6.189	3.775	-1.343
Months	85	130.235	85.607	4.000	309.000	–
R^2	85	0.094	0.148	0.001	1.000	–

Table 9: Fama-French + Carhart Four-Factor Regressions for $Obs \geq 50$.

Var	*Obs*	*Mean*	*Std*	*Min*	*Max*	*FM type t-stat*
Alpha	68	0.012	0.028	-0.034	0.186	3.450
MKTRF	68	-0.052	0.360	-2.053	0.663	-1.202
SMB	68	-0.030	0.428	-1.484	1.701	-0.578
HML	68	-0.097	0.532	-1.954	2.296	-1.504
UMD	68	-0.112	0.468	-2.465	0.696	-1.966
Months	68	156.206	74.190	50.000	309.000	–
R^2	68	0.047	0.041	0.001	0.204	–

Table 10: Fama-French + Carhart Four-Factor Regressions on Calendar-Time Portfolio Returns for March, 1981 - December, 2006 Period.

Var	$Coef$	$Std.Err.$	t-$stat$	$P > t$	95% $Confidence\ Interval$
$Alpha$	0.011	0.001	9.630	0.000	[0.009, 0.014]
$MKTRF$	-0.033	0.030	-1.110	0.266	[-0.093, 0.026]
SMB	-0.029	0.038	-0.760	0.448	[-0.103, 0.046]
HML	-0.029	0.045	-0.640	0.524	[-0.118, 0.060]
UMD	-0.107	0.027	-4.000	0.000	[-0.160, -0.055]
Obs	$F(4, 305)$	$Prob > F$	R^2	$Adj\ R^2$	$Root\ MSE$
310	4.45	0.002	0.055	0.043	0.019

Table 11: Fama-French + Carhart + Acharya_Liquidity Five-Factor Regressions on Calendar-Time Portfolio Returns for March, 1981 - December, 1999 Period.

Var	$Coef$	$Std.Err.$	t-$stat$	$P > t$	95% $Confidence\ Interval$
$Alpha$	0.013	0.001	9.420	0.000	[0.010, 0.016]
$MKTRF$	-0.016	0.036	-0.430	0.666	[-0.087, 0.055]
SMB	-0.022	0.055	-0.400	0.693	[-0.129, 0.086]
HML	-0.167	0.059	-2.810	0.005	[-0.283, -0.050]
UMD	-0.220	0.043	-5.180	0.000	[-0.304, -0.136]
$Acharya\_LIQ$	0.016	0.008	2.080	0.039	[0.001, 0.030]
Obs	$F(4, 305)$	$Prob > F$	R^2	$Adj\ R^2$	$Root\ MSE$
226	7.060	0.000	0.138	0.119	0.019

Table 12: Regression of Estimated Long-Memory Parameter ($estd$) on Governance Index (g).

Var	$Coef$	$Std.Err.$	t-$stat$	$P > t$	95% $Confidence\ Interval$
$Cons$	0.417	0.088	4.720	0.000	[0.239, 0.595]
g	0.029	0.012	2.490	0.017	[0.006, 0.052]
Obs	$F(4, 305)$	$Prob > F$	R^2	$Adj\ R^2$	$Root\ MSE$
45	6.220	0.017	0.126	0.106	0.192

Table 13: Regression of $Alpha$ on the Estimated Long-Memory Parameter ($estd$).

Var	$Coef$	$Std.Err.$	t-$stat$	$P > t$	95% $Confidence\ Interval$
$Cons$	0.049	0.009	5.380	0.000	[0.031, 0.067]
$estd$	-0.059	0.014	-4.240	0.000	[-0.087, -0.031]
Obs	$F(4, 305)$	$Prob > F$	R^2	$Adj\ R^2$	$Root\ MSE$
79	18.020	0.000	0.190	0.179	0.026

Table 14: Long-Memory Parameter and Investor Protection.

	Long – memory parameter (estd)			*Anti – self dealing index*			*Revised anti – director index*		
Country	*Mean*	*Std*	*Freq.*	*Sample fw*	*Sample ew*	*Djankov et al.*	*Sample fw*	*Sample ew*	*Djankov et al.*
Australia	0.483	0.000	1	0.757	"	"	4	"	"
Canada	0.526	0.194	39	0.642	"	"	4	"	"
South Africa	0.375	0.108	11	0.813	"	"	5	"	"
UK	0.711	0.168	15	0.950	"	"	5	"	"
US	0.606	0.188	69	0.654	"	"	3	"	"
Average English origin	**0.575**	**0.198**	**135**	**0.697**	**0.763**	**0.659**	**3.681**	**4.200**	**4.190**
Brazil	0.638	0.175	54	0.274	"	"	5	"	"
Chile	0.706	0.233	4	0.625	"	"	4	"	"
Colombia	0.845	0.152	2	0.573	"	"	3	"	"
France	0.703	0.109	6	0.379	"	"	3.5	"	"
Italy	0.784	0.118	49	0.421	"	"	2	"	"
Mexico	0.507	0.145	11	0.172	"	"	3	"	"
Portugal	0.690	0.300	3	0.444	"	"	2.5	"	"
Venezuela	0.659	0.207	4	0.092	"	"	1	"	"
Average French origin	**0.691**	**0.175**	**133**	**0.338**	**0.372**	**0.330**	**3.425**	**3.000**	**2.906**
Austria	0.495	0.334	3	0.213	"	"	2.5	"	"
Germany	0.646	0.154	51	0.282	"	"	3.5	"	"
South Korea	0.714	0.123	57	0.469	"	"	4.5	"	"
Switzerland	0.689	0.207	10	0.267	"	"	3	"	"
Average German origin	**0.678**	**0.155**	**121**	**0.367**	**0.307**	**0.383**	**3.905**	**3.375**	**3.036**
Denmark	0.659	0.152	21	0.463	"	"	4	"	"
Finland	0.721	0.187	14	0.457	"	"	3.5	"	"
Norway	0.666	0.123	8	0.421	"	"	3.5	"	"
Sweden	0.557	0.196	48	0.333	"	"	3.5	"	"
Average Scandinavian origin	**0.615**	**0.189**	**91**	**0.390**	**0.418**	**0.386**	**3.615**	**3.625**	**3.800**
Average Civil Law	**0.666**	**0.175**	**345**	**0.362**	**0.368**	**0.350**	**3.643**	**3.250**	**3.029**
World Average	**0.640**	**0.186**	**480**	**0.456**	**0.462**	**0.440**	**3.654**	**3.476**	**3.368**

Table 15: Long-Memory Parameter and Investor Protection (Mean Comparison).

	T-Stat	
	Long − memory parameter (estd) Sample	*Revised Anti − director index Djankov et al.*
Common vs. Civil	-4.713	4.46
French vs. Common	5.091	-4.49
French vs. German	0.634	-0.37
French vs. Scandinavian	3.036	-1.73

	T-Stat – Significance Level (single sided)	
	Long − memory parameter (estd) Sample	*RevisedAnti − director index Djankov et al.*
Common vs. Civil	0.0%	0%
French vs. Common	0.0%	0%
French vs. German	26.3%	71%
French vs. Scandinavian	0.1%	9%