COMOVEMENT AND MOMENTUM

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Abstract

Recent evidence of excessive comovement among stocks following index additions (Barberis, Shleifer, and Wurgler, 2005) and stock splits (Green and Hwang, 2009) challenges traditional finance theory. Based on a simple model, we show that the bivariate regressions relied upon in the literature often provide little or no information about the economic magnitude of the phenomenon of interest. Instead, univariate regressions of the stock return on the returns of the group it is leaving (e.g., non-S&P stocks) and the group it is joining (e.g., S&P stocks) reveal relevant information. Moreover, results in the literature are consistent with changes in the fundamental factor loadings of the stocks. When we reexamine the empirical evidence using control samples matched on past returns and compute Dimson betas, almost all evidence of excess comovement disappears. One key element to understanding these striking results is that, in both the examples we study, the stocks exhibit strong returns prior to the event in question. Relatedly, we document the heretofore unknown empirical regularity that winner stocks exhibit increases in betas. Thus, the apparent excess comovement is just a manifestation of momentum.

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1. Introduction

In a perfect and frictionless financial market, asset prices change to reflect new information about future cash flows and discount rates. To the extent that there are common factors affecting either cash flow or discount rates, asset prices would move together to reflect innovations in such common factors.

However, there is growing evidence that prices move together for reasons that are seemingly unrelated to fundamentals. Evidence of this excess comovement has been found among S&P 500 index additions and deletions (Vijh, 1994; Barberis, Shleifer, and Wurgler, 2005), changes in S&P 500 value and growth indices (Boyer, 2011), changes in the Nikkei 225 index (Greenwood and Sosner, 2007), changes in the UK indices (Mase, 2008), changes in Nikkei 225 index weights (Greenwood, 2008), additions to many national market indices (Claessens and Yafeh, 2011), stock splits (Green and Hwang, 2009), stocks with correlated trading among retail investors (Kumar and Lee, 2006), stocks with corporate headquarters in the same geographic area (Pirinsky and Wang, 2006), institutional ownership (Pindyck and Rotemberg, 1993), stocks in closed-end country funds (Hardouvelis et al., 1994; Bodurtha et al., 1995), stocks in closed-end domestic funds (Lee et al., 1991), sovereign bonds (Rigobon, 2002), and commodity futures (Tang and Xiong, 2012).

Though excessive comovement in stock returns is attributed to several non-fundamental factors, the primary explanation is an asset class effect, which is created by correlated demand unrelated to fundamentals for assets in a particular class. Theoretical models developed by Basak and Pavlova (2013), DeMarzo, Kaniel and Kremer (2004), and Barberis and Shleifer (Barberis, Shleifer, and Wurgler, 2005) propose three sources of friction and investor sentiment. Excess investor demand for a particular group of securities may arise because of investor awareness (habitat) or because those stocks form an asset class that is easy to follow (category). Third, the speed of information diffusion may increase for stocks included in the index. Similar arguments are in Hou and Moskowitz (2005) and Pindyck and Rotemberg (1993). Improvement in price discovery would cause the added stock to comove more strongly with index stocks than with non-index stocks. Since it is difficult to empirically distinguish between the first two views, Greenwood (2008) combines them into a single demand-based theory, or an asset class effect. The last source of friction, quicker adjustment in prices to new information is a desirable outcome of index additions because it makes prices more efficient even though it may increase comovement. In other words, there was too little comovement in the absence of efficient information diffusion, which has now been increased to an appropriate level (Claessens and Yafeh, 2011). Other explanations relate to transactions costs at an index level versus an individual stock level. However, we focus on the asset class effect as the generally accepted source of comovement.
(2003), among others, are consistent with such an asset class effect. However, the sources of this correlated demand are varied: investor behavior that causes investors to choose stocks based on styles or categories (Barberis and Shleifer, 2003); agents who care about relative wealth choosing assets held by other members of the community (DeMarzo, Kaniel and Kremer, 2004); or institutional investors who care about their performance relative to an index tilting their portfolios towards stocks that are in that index (Basak and Pavlova, 2013).

Two recent papers, von Drathan (2013) and Kasch and Sarkar (2013), challenge the empirical evidence mentioned above in the context of two specific events, FTSE 100 and S&P500 index turnover, respectively. They both point out that these events coincide with changes in fundamentals. Our focus is on providing a more general view of the issue and on understanding the mechanisms that underlie the regression results in the existing literature, as explained below.

In this paper, we reexamine the evidence on comovement, focusing on two recent studies that document what appears to be strong support for this phenomenon, but in apparently unrelated contexts. The first is Barberis, Shleifer, and Wurgler (2005), which is considered a classic paper on comovement. Their sample consists of stocks that enter or leave the S&P500, an event that has been used by many other studies because index changes are generally believed to have little fundamental effect on the firm being added to or deleted from the index (Chen et al., 2004; Elliott et al., 2006). Their hypothesis is that stocks in the index comove more with index stocks, whereas those not in the index comove more with non-index stocks. The second paper is Green and Hwang (2009), who study comovement before and after stock splits. Specifically, their argument is that stocks with similar price levels comove more than would be justified by fundamentals, i.e., that a stock moves more with high-priced stocks prior to a split and more with low-priced stocks after a split. As with index changes, splits appear to be useful events to study because they do not affect splitting firms in any fundamental way, although the announcement may signal private information.

In both cases, the primary evidence is in the form of differences between the coefficients of two regressions conducted before and after the event: (1) a univariate regression of the stock return on the return of the group it is joining, and (2) a bivariate
regression of the stock return on both the old group and the new group. The bivariate regression results in Barberis, Shleifer, and Wurgler (2005) show that for additions to the S&P 500 index, their coefficient on S&P 500 returns increases dramatically after they join the index while the coefficient on non-index stocks declines. In a similar vein, the bivariate regression results in Greene and Hwang (2009) show that stocks after a split load more heavily on low-priced stocks (the new group) and less on high-priced stocks (the old group).

In order to better understand the implications of the excess comovement hypothesis for stock returns, we first develop a model closely related to that of Barberis, Shleifer, and Wurgler (2005). Some implications of our model are similar to those derived in their paper, but we highlight three additional important implications. First, the model suggests that a univariate regression of the stock return on the return of the old group after the event can be very informative, a specification not examined in Barberis, Shleifer, and Wurgler (2005) or Green and Hwang (2009).

Second, the model indicates that the results of the bivariate regressions estimated by Barberis, Shleifer, and Wurgler (2005) and Green and Hwang (2009) are extremely sensitive to small changes in the parameters. The sensitivity of these types of regression coefficients has been documented in the literature (Spanos and McGuirk, 2002). Most critically for our analysis, this sensitivity implies that the interpretation of these coefficient estimates is not straightforward, and that they may well provide little or no information about the question of economic interest—how much, if at all, is excess comovement responsible for the variation in stock returns.

Third, the model shows that changes in the parameters around the events, in particular shifts in loadings on the fundamental factor, can affect the univariate regression results. For example, an increase in the beta of a stock in the sample will generate an increase in the coefficient of the stock on the new group return after the event. In other words, these empirical results are also consistent with a change in fundamental comovement not just excess comovement. Of course, this phenomenon also has implications for the univariate regression of the stock return on the old group return discussed above, and, in fact, it is this regression that allows us to distinguish between the two competing explanations.
We begin our empirical analysis by reexamining comovement following index changes. We expand the Barberis, Shleifer, and Wurgler (2005) sample period of 1976-2000 to 1962-2011. Using daily data, where Barberis, Shleifer, and Wurgler (2005) report their strongest results, there is no evidence of excess comovement in the 1962-76 period, which is consistent with the excess comovement story because indexing was in its infancy during that period with less than 1% of the assets indexed to the S&P 500. As in Barberis, Shleifer, and Wurgler (2005), the bivariate regression results show a significant increase in beta relative to the S&P 500 index and a significant decrease in beta relative to the old index. We address limitations of the bivariate regressions in greater detail later. In general, based on the two univariate regressions, we find that stocks added to the S&P 500 index move more with the S&P 500 index but they also move more with the old group of non-S&P index stocks. The difference in beta changes is not significant for the 1976-87 period, nor is it significant for the 2001-11 period. The difference in beta changes is, however, significant for the 1988-2000 period.

For the stock split sample, the bivariate regressions again show an increase in the beta with the new group, though there is no statistically significant decrease in beta relative to the old group. Evidence in support of comovement is much weaker when the univariate regressions are analyzed: the increase in beta between returns on splitting stocks and returns on the new group (i.e., low-priced stocks) is almost equal to the increase in beta between returns on splitting stocks and returns on the old group (i.e., high-priced stocks).

These initial empirical results indicate that it may be increases in the fundamental betas of the stocks around the events that are driving much of the results reported in the literature as excess comovement. The natural question is why do these betas increase, i.e., what do stocks added to the S&P500 and those undergoing splits have in common? The answer is that both groups of stocks exhibit positive performance prior to the event. In the language of the literature on cross-sectional momentum effects, they are winners. Following the usual momentum methodology, we find that betas of winner stocks increase during the formation period and continue to increase during the holding period, before declining at longer horizons. Therefore, it is likely that at least some of the results reported by Barberis, Shleifer, and
Wurgler (2005) and Green and Hwang (2009) are caused by the inclusion of momentum stocks in their samples.

Given the apparent importance of fundamental betas, we next turn to a more refined analysis that attempts to better measure and control for these beta changes. First, we improve the estimation of the betas by employing a Dimson (1979) approach to adjust for non-synchronous trading using leads and lags of the relevant indices in the regressions. Though the S&P 500 index consists of some of the largest stocks in the U.S. economy, index changes are concentrated mainly among the smaller stocks in the index. The relatively smaller size of index additions implies that those stocks may not trade as often as other stocks in the index, but they may trade more frequently than non-index stocks, contributing to the higher correlation with S&P 500 stocks relative to non-S&P 500 stocks. Similarly, the trading frequency of stocks that split may differ from that of the stocks in either the low- or high-priced indices that we construct. We add two leads and lags of the index returns to pick up these effects.

Second, the impact of momentum and the associated beta changes on measured comovement is estimated using a matched sample approach. For each index change and stock split, we choose a firm in the same size decile that comes closest based on momentum, i.e., has a similar return over the past year. We then adopt a difference in difference in difference approach, examining the differences in the changes of the betas before and after the event across the stocks in the original sample and the matched sample.

The empirical results from this refined analysis are striking. For both S&P500 index additions and stock splits, the original sample and matched sample stocks exhibit differences in beta changes that are not significantly different. In other words, the differences between the changes across the two univariate regressions are statistically indistinguishable for the sample and control stocks. This result is compelling evidence that the apparent excess comovement is actually driven by changes in loadings on the fundamental component of returns, not by asset class effects. The control stocks also show similar changes in bivariate regression coefficients before and after the event to which the sample stocks are subject. Moreover, this result is not simply an artifact of limited statistical power. The point estimates indicate that excess comovement is not economically significant either.
A breakdown of our two adjustments, i.e., the Dimson adjustment and the matched control adjustment, shows that their importance differs dramatically for the two samples. For the stock split sample, the Dimson adjustment does little, but the momentum control is critical because these stocks exhibit very strong past performance and resulting beta changes. In contrast, for the S&P500 index addition sample, the momentum effect is somewhat weaker and both adjustments are necessary. The differential momentum effect is consistent with a significantly greater proportion of winner stocks that split than the proportion of winner stocks that are added to the S&P500 index.

The paper is organized as follows. In the next section, we introduce the model and examine its implications for univariate and bivariate regressions. Section 3 describes the data and methodology for momentum, index changes, and stock splits. Section 4 contains the main empirical results for the original sample. In Section 5, we examine the link between momentum and beta changes and then reexamine the data in the light of this evidence. We perform several robustness checks in Section 6, and Section 7 concludes.

2. A Model

In order to understand the implications of the regression results reported in the literature for the economic importance of the excess comovement phenomenon, it is useful to write down a relatively simple and stylized model in which the coefficients in these regressions can be calculated in closed form. Our goal is not to fully capture reality, but rather, in the spirit of the model in Barberis, Shleifer, and Wurgler (2005), to generate some general insights and predictions that we can use to interpret the subsequent empirical results. Our model is not identical to that in Barberis, Shleifer, and Wurgler (2005), although the key predictions are similar, because we want to construct the simplest possible model that both highlights the features of the univariate and bivariate regressions that we believe are important and captures the essence of the excess comovement hypothesis.

2.1 Setup and Assumptions
Denote as $y_t$ the return on a stock that is changing membership between groups 1 and 2 with returns $x_{1t}$ and $x_{2t}$, respectively, e.g., non-S&P and S&P stocks or high-priced and low-priced stocks:

$$
y_t = b_{yt}f_t + c_{1t}u_{1t} + c_{2t}u_{2t} + e_{yt}
$$

$$
x_{1t} = b_{1t}f_t + u_{1t} + e_{1t}
$$

$$
x_{2t} = b_{2t}f_t + u_{2t} + e_{2t}
$$

$$
\text{var}(e_{yt}) \equiv \sigma^2_{eit} \quad \text{var}(u_{1t}) \equiv \sigma^2_{uit} \quad \text{var}(f_t) \equiv \sigma^2_{f_t}
$$

where $f$ is the fundamental, common return shock, which could easily be extended to a multi-factor context; $u_i$ are group-specific, non-fundamental return shocks; and $e_i$ are idiosyncratic fundamental return shocks.

For identification purposes assume

$$
\text{cov}(u_{1t}, u_{2t}) = 0
$$

$$
\text{cov}(u_{1t}, f_t) = \text{cov}(e_{1t}, f_t) = 0 \quad \forall i, j
$$

$$
\text{cov}(u_{i1}, e_{jt}) = 0 \quad \forall i, j
$$

That is, non-fundamental, group-specific shocks are assumed to be uncorrelated across groups; the common fundamental factor is uncorrelated with the other shocks; and the idiosyncratic, fundamental shocks are uncorrelated with the non-fundamental shocks.

The economic content of the excess comovement hypothesis is a statement about the loadings of stock $y$ on the two non-fundamental, group-specific shocks, $u_1$ and $u_2$. Specifically, using underbars and overbars to denote values prior to and after the stock switches from group 1 to group 2, the theoretical predictions of this hypothesis are

$$
\bar{c}_{1t} = \bar{c}_1 > 0 \quad \bar{c}_{2t} = 0
$$

$$
\bar{0}_{1t} = 0 \quad \bar{0}_{2t} = \bar{c}_2 > 0
$$

i.e., there is a zero loading on the group-specific shock of the group to which the stock does not belong, and a positive loading on the group-specific shock of the group to which the stock does belong. We also assume that all the other parameters of the model are constant in each sub-period, i.e., the periods before and after the move of stock $y$ between the groups, but that they
can vary across the sub-periods. As above, we use underbars and overbars to designate these parameters.

A natural measure of excess comovement is the percentage of the variation in stock $y$’s return that is due to this excess comovement, both prior to and after the event:

$$\frac{c_{1}^{2} \sigma_{u1}^{2}}{\sigma_{y}^{2}} \quad \text{and} \quad \frac{c_{2}^{2} \sigma_{u2}^{2}}{\sigma_{y}^{2}}$$ (4)

This measure is equivalent to the R-squared one would get if one regressed the stock return on the non-fundamental component of the corresponding group return.

2.2 Implications

Consider the following three regressions run pre- and post-switch:

$$y_{t} = \alpha + \beta_{1} x_{1r} + \epsilon_{t}$$
$$y_{t} = \alpha + \beta_{2} x_{2r} + \epsilon_{t}$$
$$y_{t} = \alpha + \beta_{1b} x_{1r} + \beta_{2b} x_{2r} + \epsilon_{t}$$ (5)

The probability limits of the univariate regression coefficients under the model above are

$$\beta_{1} = \frac{b_{1} \bar{b}_{1} \sigma_{f}^{2} + c_{1} \sigma_{u1}^{2}}{\sigma_{x1}^{2}} \quad \bar{\beta}_{1} = \frac{\bar{b}_{1} \bar{b}_{2} \sigma_{f}^{2}}{\sigma_{x1}^{2}}$$
$$\sigma_{x1}^{2} = b_{1}^{2} \sigma_{f}^{2} + \sigma_{u1}^{2} + \sigma_{e1}^{2} \quad \bar{\sigma}_{x1}^{2} = \bar{b}_{1}^{2} \sigma_{f}^{2} + \sigma_{u1}^{2} + \sigma_{e1}^{2}$$

$$\beta_{2} = \frac{b_{2} \bar{b}_{2} \sigma_{u2}^{2}}{\sigma_{x2}^{2}} \quad \bar{\beta}_{2} = \frac{\bar{b}_{2} \bar{b}_{2} \sigma_{u2}^{2}}{\sigma_{x2}^{2}}$$
$$\sigma_{x2}^{2} = b_{2}^{2} \sigma_{f}^{2} + \sigma_{u2}^{2} + \sigma_{e2}^{2} \quad \bar{\sigma}_{x2}^{2} = \bar{b}_{2}^{2} \sigma_{f}^{2} + \sigma_{u2}^{2} + \sigma_{e2}^{2}$$ (6)

For the bivariate regression
\[ \beta_{1b} = \frac{1}{1 - \rho_{x_1, x_2}^2} \left[ \beta_1 - \rho_{x_1, x_2}^2 \frac{\sigma_{x_1}^2}{\sigma_{x_1}^2} \beta_2 \right] \quad \beta_{2b} = \frac{1}{1 - \rho_{x_1, x_2}^2} \left[ \beta_2 - \rho_{x_1, x_2}^2 \frac{\sigma_{x_1}^2}{\sigma_{x_1}^2} \beta_1 \right] \]
\[ \bar{\beta}_{1b} = \frac{1}{1 - \rho_{x_1, x_2}^2} \left[ \bar{\beta}_1 - \rho_{x_1, x_2}^2 \frac{\bar{\sigma}_{x_1}^2}{\bar{\sigma}_{x_1}^2} \bar{\beta}_2 \right] \quad \bar{\beta}_{2b} = \frac{1}{1 - \rho_{x_1, x_2}^2} \left[ \bar{\beta}_2 - \rho_{x_1, x_2}^2 \frac{\bar{\sigma}_{x_1}^2}{\bar{\sigma}_{x_1}^2} \bar{\beta}_1 \right] \] (7)
\[ \rho_{x_1, x_2} = \frac{\text{cov}(x_1, x_2)}{\sigma_{x_1} \sigma_{x_2}} \quad \text{cov}(x_1, x_2) = b_1 b_2 \sigma_f^2 + \text{cov}(\varepsilon_1, \varepsilon_2) \]
\[ \bar{\rho}_{x_1, x_2} = \frac{\text{cov}(\bar{x}_1, \bar{x}_2)}{\bar{\sigma}_{x_1} \bar{\sigma}_{x_2}} \quad \text{cov}(\bar{x}_1, \bar{x}_2) = \bar{b}_1 \bar{b}_2 \bar{\sigma}_f^2 + \text{cov}(\bar{\varepsilon}_1, \bar{\varepsilon}_2) \]

(see the appendix for detailed derivations).

Furthermore, if the covariance matrix of return shocks is constant,

\[ b_j = \bar{b}_j \equiv b_j \quad \sigma_f^2 = \bar{\sigma}_f^2 \equiv \sigma_f^2 \quad \sigma_{u_1}^2 = \bar{\sigma}_{u_1}^2 \equiv \sigma_{u_1}^2 \quad \sigma_{e_i}^2 = \bar{\sigma}_{e_i}^2 \equiv \sigma_{e_i}^2 \] (8)
i.e., the weights on the common factor, the variances of the non-fundamental shocks, and the variances of the fundamental shocks do not change from pre- to post-switch, then

\[ \beta_1 > \bar{\beta}_1 \quad \beta_2 > \bar{\beta}_2 \quad \beta_{1b} > \bar{\beta}_{1b} \quad \beta_{2b} > \bar{\beta}_{2b} \] (9)

(again, see the appendix for details). Intuitively, when the stock switches from group 1 to group 2, it begins to move with the non-fundamental shock to group 2 and ceases to move with the non-fundamental shock to group 1; therefore, its coefficient on group 1 returns decreases and its coefficient on group 2 returns increases, both in a univariate and a bivariate context.

If we further assume that (i) the groups are fundamentally well-diversified, i.e., there is no idiosyncratic fundamental shock at the group level (\( \sigma_{e_1}^2 = \sigma_{e_2}^2 = 0 \)), (ii) stock \( y \) has a weight of one on the non-fundamental group shock, i.e., \( c_1 = \bar{c}_2 = 1 \), and (iii) the loadings on the fundamental shocks are all equal to unity, i.e., \( b_y = b_1 = b_2 = 1 \), then we duplicate the more specific results contained in Prediction 2 of Barberis, Shleifer, and Wurgler (2005):\(^\text{2}\)

\(^\text{2}\) See the appendix for details. This result is not identical to that in Barberis, Shleifer, and Wurgler (2005). Specifically, their result is slightly weaker: \( \beta_{1b} = 1, \beta_{2b} = 0 \quad 0 < \beta_{1b} < 1, 0 < \beta_{2b} < 1, \beta_{1b} + \beta_{2b} = 1 \)
This result is important because it illustrates a flaw in the interpretation of the bivariate regression coefficients. From an economic standpoint, we are not directly interested in these coefficients; the key parameters are the loadings of the stock return on the various factors in equation (1) and the variances of these factors, which determine the measure of excess comovement defined in equation (4) above. However, under the assumptions outlined above, the bivariate regression coefficients are completely independent of the variances of the non-fundamental component of group and stock returns as long as these quantities are strictly positive. Thus, even when the non-fundamental component of both stock and group returns is economically meaningless, in the sense that it contributes essentially nothing to the variability of returns, the bivariate coefficients appear to suggest a dramatic and economically meaningful change in the comovement properties of returns.

Of course, this extreme invariance result does depend on the assumed factor loadings, specifically the fact that the stock and the groups load equally on both the fundamental and non-fundamental factors. However, in more general settings, it is still the case that the coefficients in the bivariate regression are sensitive to small changes in the parameters of the driving processes, and their magnitudes do not reflect the quantities of economic interest. The intuition is that all reasonably well-diversified stock portfolios tend to be very highly correlated. Thus, the correlation between the returns on the two groups of stocks will be close to one. This issue is the multi-collinearity in the bivariate regression that is discussed in Barberis, Shleifer, and Wurgler (2005). As they rightly point out, multi-collinearity does not affect the consistency of the estimates in OLS. But, as the example above illustrates, the coefficients in the bivariate regression may tell us very little, or even nothing, about what we really want to know, i.e., how much excess comovement affects returns. This concern is especially relevant if the strong

\[ \beta_{1b} = 1, \beta_{2b} = 0, \beta_{1b} = 0, \beta_{2b} = 1 \] (10)
assumptions above about the stability of the parameters across the two sub-periods, which are critical in deriving the results, are not valid.

Fortunately, the coefficients in the univariate regressions isolate precisely the quantities of interest. Going back to the more general assumptions about stability of the parameters across the sub-periods, but making no assumptions about the magnitudes of the factor loadings, the differences between these coefficients pre- and post-switch are (see the appendix for details):

\[
\begin{align*}
\Bar{\beta}_1 - \beta_1 &= -\frac{c_1\sigma_{u_1}^2}{\sigma_{x_1}^2} \\
\Bar{\beta}_2 - \beta_2 &= \frac{c_2\sigma_{u_2}^2}{\sigma_{x_2}^2}
\end{align*}
\]

(11)

Thus, empirical evidence that the coefficient on the return of the group to which a stock is moving (group 2) increases after the switch would appear to be strong evidence of excess comovement, and the magnitude of this difference multiplied by the variance of the group 2 stock return would measure the economic importance of this phenomenon.\(^4\)

However, there are two cautionary notes. First, the excess comovement hypothesis would suggest a decrease in the coefficient on the return of the group from which a stock is moving (group 1). Second, these results do depend on the assumption about the stability of the other parameters across the sub-periods.

It is worthwhile examining this latter issue in more detail. Consider, for example, the case where the loading of stock \(y\) on the fundamental factor, \(b_{y,t}\), is allowed to vary across the sub-periods. In this case

\[
\begin{align*}
\Bar{\beta}_1 - \beta_1 &= \frac{(\Bar{b}_y - b_y)\sigma_{b_{y}}^2 - c_1\sigma_{u_1}^2}{\sigma_{x_1}^2} \\
\Bar{\beta}_2 - \beta_2 &= \frac{(\Bar{b}_y - b_y)\sigma_{b_{y}}^2 + c_2\sigma_{u_2}^2}{\sigma_{x_2}^2}
\end{align*}
\]

(12)

\(^4\)This quantity, \(\sigma_{x_2}^2(\Bar{\beta}_2 - \beta_2)\), is the difference between the covariances of the stock return and the group return in the two sub-periods, i.e., \(\text{cov}(\Bar{y}_t, \Bar{x}_{2t}) - \text{cov}(\Bar{y}_t, \Bar{x}_{1t})\).
That is, an increase in the coefficient on the group 2 return could be attributable in part, or completely, to an increase in stock \( y \)'s loading on the fundamental factor. However, the difference between the weighted coefficient changes still measures the magnitude of the effects of excess comovement if the loadings of the two groups on the fundamental factor are the same (i.e., \( b_1 = b_2 \)):

\[
\sigma_x^2 (\bar{\beta}_2 - \bar{\beta}_1) - \sigma_x^2 (\bar{\beta}_{11} - \bar{\beta}_{12}) = c_2 \sigma_{u2}^2 + c_1 \sigma_{x1}^2
\]  

(13)

While permitting changes in the fundamental loading of the stock, this result relies on (approximate) equality of the fundamental loadings of the two groups. One way to get around this assumptions is to consider a matched stock (denoted \( m \)) that starts in the same group as stock \( y \), has similar characteristics across the relevant dimensions, but does not switch from group 1 to group 2. For this stock, the univariate regression coefficients are

\[
\hat{\beta}_{m1} = \frac{b_{m1} \sigma_{f1}^2 + c_{1} \sigma_{m1}^2}{\sigma_{x1}^2} \quad \bar{\beta}_{m1} = \frac{b_{m1} \sigma_{f}^2 + c_{1} \sigma_{m1}^2}{\sigma_{x1}^2}
\]

\[
\hat{\beta}_{m2} = \frac{b_{m2} \sigma_{f2}^2}{\sigma_{x2}^2} \quad \bar{\beta}_{m2} = \frac{b_{m2} \sigma_{f}^2}{\sigma_{x2}^2}
\]

(14)

again maintaining the parameter stability assumptions across the sub-periods for the group returns. Intuitively, this matched stock continues to load on the group 1 non-fundamental factor after the event but never loads on the group 2 non-fundamental factor. Thus, taking differences in differences in differences (see the appendix for details)

\[
(\Delta \beta_2 - \Delta \beta_{m2}) - (\Delta \beta_1 - \Delta \beta_{m1}) = (\Delta b_y - \Delta b_m \left( \frac{b_{y} \sigma_{f}^2}{\sigma_{x2}^2} - b_{m} \sigma_{f}^2 \sigma_{x1}^2 \right)) + \left( c_2 \frac{\sigma_{u2}^2}{\sigma_{x2}^2} + c_1 \frac{\sigma_{m1}^2}{\sigma_{x1}^2} \right)
\]

\[
\Delta \beta_1 \equiv \bar{\beta}_1 - \beta_1 \quad \Delta \beta_2 \equiv \bar{\beta}_2 - \beta_2 \quad \Delta \beta_{m1} \equiv \bar{\beta}_{m1} - \beta_{m1} \quad \Delta \beta_{m2} \equiv \bar{\beta}_{m2} - \beta_{m2}
\]

\[
\Delta b_y \equiv \bar{b}_y - \beta_2 \quad \Delta b_m \equiv \bar{b}_m - \beta_m
\]

(15)

As long as the matched firms experience the same change in loading on the fundamental factor as stock \( y \) (i.e., \( \Delta b_y - \Delta b_m = 0 \)), we can recover an equivalent measure of excess comovement under weaker assumptions.
The fundamental conclusion from this stylized model is that it is the univariate coefficients from the regressions of stock $y$’s return on the two group returns, or more precisely the covariances between the stock return and the group returns, only one of which is examined in the literature, that are informative about the excess comovement hypothesis. Moreover, the coefficients from a bivariate regression are likely misleading and possibly completely irrelevant for assessing the economic importance of this phenomenon.

2.3 Numerical Examples

A few simple numerical examples serve to illustrate the main points made analytically in Section 2.2. Consider the simplified model for stock $y$ returns before and after the event:

$$y_t = b_y f_t + c_1 u_{1t} + e_{yt}$$
$$y_t = b_y f_t + c_2 u_{2t} + e_{yt}$$

and after the event. Group returns and variances are assumed stable across the event:

$$x_{1t} = b_1 f_t + u_{1t} + e_{1t}$$
$$x_{2t} = b_2 f_t + u_{2t} + e_{2t}$$

$$\text{var}(e_{it}) \equiv \sigma_{ei}^2 \quad \text{var}(u_{it}) \equiv \sigma_{ui}^2 \quad \text{var}(f_t) \equiv \sigma_f^2$$

Example 1: $\bar{b}_y = b_y = b_1 = b_2 \quad c_1 = c_2 = 1 \quad \sigma_{ei}^2 = \sigma_{i1}^2 = 0$

All fundamental factor loadings are equal, the non-fundamental factor loadings equal unity, and the groups are well-diversified, i.e., they have no fundamental idiosyncratic return component. In this case, regardless of the economic magnitude of excess comovement, i.e., regardless of the magnitudes of the variances of the non-fundamental components of the group and stock returns, $\sigma_{ui}^2$, the bivariate regression coefficients are constant and equal to one or zero:

$$\beta_{1b} = 1, \beta_{2b} = 0 \quad \overline{\beta}_{1b} = 0, \overline{\beta}_{2b} = 1$$

This is the result given in equation (10) and derived in the appendix.
3. Data and Methodology

The CRSP stock files at the University of Chicago and Standard and Poor’s are the primary sources of data. In general, we follow the methodologies in Barberis, Shleifer, and Wurgler (2005) and Green and Hwang (2009) for constructing our tests. For index changes, we follow the methodology of Barberis, Shleifer, and Wurgler (2005) except that we use only daily data because their results are weaker with weekly and monthly data. Barberis, Shleifer, and Wurgler (2005) use additions to the S&P 500 from 1976 to 2000 and deletions from 1979 to 2000, whereas our initial sample extends from 1962 to 2011 for index additions.\(^5\) However, subperiod analysis corresponds to their subperiods. Index deletions are evaluated for robustness in Section 6. Like Barberis, Shleifer, and Wurgler (2005), we estimate betas in the pre-inclusion period using 12 months of data ending the month before the announcement of the stocks addition to the S&P 500 and betas in the post-inclusion period using 12 months of data starting the month after the inclusion of the stock in the S&P 500.

For stock splits, we follow the methodology in Green and Hwang (2009) and the clarifications obtained directly from the authors, though some differences in methodology persist. Like Green and Hwang (2009), our sample consists of all common stocks where the stock price was $10 or more before the stock split.\(^6\) The high-price index consists of stocks whose prices are ±25% of the price of the splitting stock just prior to the split. The low price index consists of all stocks whose price is above $5 and within ±25% of the post-split price calculated based on the pre-split price and the split ratio.

For momentum, we follow a methodology that is similar to that in Jegadeesh and Titman (2001) and form momentum portfolios using a 12-month formation period, one skip month, and 12-month holding period. More specifically, at the end of each June from 1962 through 2010, stocks with a price of at least $10 that do not fall into the bottom size decile of NYSE stocks are assigned to 10 momentum deciles based on their cumulative returns over the

\(^5\) We limit our main analysis to index additions with share codes of 10 and 11 to remain consistent with Barberis, Shleifer, and Wurgler (2005). However, the results are similar if the sample contains all index additions.

\(^6\) For consistency with their results, we only include stock splits identified by CRSP with a distribution code ‘5523’. However, inclusion of stocks splits with a CRSP distribution code ‘5533’ produces similar results.
preceding 252 days.\textsuperscript{7} We estimate betas for each stock based on a rolling window of 252 days from two years before formation of momentum portfolios through two years after formation, and compare beta changes for the top and bottom momentum portfolios. Thus, betas for years -2 and -1 are estimated over rolling windows ending 504 and 252 trading days before portfolio formation, respectively. Post-formation momentum portfolio betas allow for a 21-trading day skip, and are estimated over 252 days ending 273 and 525 trading days after portfolio formation. The top return decile and the bottom return decile in the formation period are identified as winner stocks and loser stocks respectively.

4. Empirical Results: Univariate and Bivariate Regressions

The first step in our analysis is to recreate and reexamine the univariate and bivariate regressions reported in the literature for the S&P500 index addition and stock split samples. These are the regressions specified in equation (5), and they are estimated twice, once before the event and once after. Note that the first regression, the return on the stock on the return of the group that it is leaving, is not examined in the literature. The implications of the coefficients in these regressions for the excess comovement hypothesis are discussed in Section 2.2.

The results are presented in Tables 1 and 2 for S&P500 index additions and stock splits, respectively. In each case, Panel A shows the univariate regression results and Panel B the bivariate regression results. In Panel A, the set of 4 columns beginning with the third column contain the betas relative to the old group portfolio (non-index stocks or high-priced stocks) before and after the event and the associated changes, the next set of 4 columns contain the analogous numbers relative to the new group portfolio, and the final 3 columns show various measures of the difference between the changes of coefficients across the event. In Panel B, the sets of columns present the coefficients in the bivariate regression, their changes, the differences between beta changes for the two groups, and R-squareds. Where possible, we

\textsuperscript{7} The sample ends in 2010 because we are evaluating beta changes up to two years after formation of momentum portfolios.
report results for the full period and sub-periods corresponding to those used in the original study.

Turning first to the S&P500 additions sample, the results from the univariate regressions on the S&P500 index (the new group) for two sub-periods, 1976-87 and 1988-2000 are consistent with those reported by Barberis, Shleifer, and Wurgler (2005) in their Panel A of Table 1. For 1976-87, we report a change in beta of 0.062 ($\Delta \beta_2$) based on a sample of 197 index additions compared with 0.067 in Barberis, Shleifer, and Wurgler (2005) based on a sample of 196 index additions. For 1988-2000, we and Barberis, Shleifer, and Wurgler (2005) both find an increase in beta of 0.214 after stocks are added to the S&P 500 index. Looking at all the periods analyzed in this study, we note that there is no effect in the first sub-period (1962-76), when indexing was almost non-existent, but the effect becomes increasingly positive through the year 2000. Interestingly, however, this difference is less than a third as large (0.064 vs. 0.214) for the very last sub-period, 2001-2011, which was not covered in the original sample. Notwithstanding this anomaly, on their own, these results would naturally be interpreted, in the context of the model in Section 2, as evidence of excess comovement. The stock begins to load more heavily on the index return after it joins the index.

Looking at the univariate results with the non-index returns as the independent variable shows that this simple interpretation cannot be completely accurate. To be consistent with excess movement, the change in the coefficient relative to the old group from before to after the stock joins the index ($\Delta \beta_1$) should be negative. That is, the stock should load less heavily on non-index returns when it is not in the index, a change not examined in prior studies. Instead, we find that this change ($\Delta \beta_1$) is approximately equal in magnitude to the coefficient change for the other regression ($\Delta \beta_2$) for the 1976-87 and 2001-11 periods. Consequently, the measure of total excess comovement, the difference between these changes ($\Delta \beta_2 - \Delta \beta_1$), is small and statistically insignificant for these two subperiods. Taken together, these results suggest that it

---

8 Standard and Poor’s did not publicly announce index changes until September 1976. Therefore, the first period for index changes runs from 1962 to August 1976, whereas the second period begins in September 1976. However, for ease of reference, we term the periods 1962-76 and 1976-87.

9 We evaluate the 1962-76 period more fully in the Robustness Checks section.
may be changes in loadings on the fundamental factor that are more important, except for the 1988-2000 sub-period.

The model in Section 2 implies that the bivariate results are unreliable in terms of assessing the economic magnitude of any excess comovement, but, for completeness, we present results from the bivariate regressions in Panel B. These results are similar to those reported in Barberis, Shleifer, and Wurgler (2005) for matching subperiods. Their bivariate regressions show an increase in the beta with the S&P index (new index) and a decrease in the beta with a non-S&P 500 index (old index). We return to the bivariate results in Section 5.

The results for stock splits are reported in Table 2, with the Green and Hwang (2009) sub-periods. The changes in beta relative to the new index reported in Table 2 for matching sub-periods are very close to those reported by Green and Hwang (2009) in their Panel A of Table 2: we report a of 0.196 for 1971-1990 with a sample of 2,349 splits compared to their change in beta of 0.204 with a sample of 2,302 splits for the same period. For the 1991-2004 period, the samples are marginally different: Green and Hwang (2009) report an increase of 0.255 in beta with a sample of 2,303 splits compared to 0.248 with a sample of 2,478 splits in this paper. As for index changes, the univariate regressions results are striking. The coefficient on low-priced stocks increases significantly after the split for all sub-periods and is consistent with the notion of excess comovement documented in the earlier studies.

We also examine the change in beta relative to the old, high-priced index before and after the split. From Panel A of Table 2, we can see that $\Delta \beta_1$ is significantly positive for all sub-periods, which suggests that the beta of the splitting stock increases not only relative to the new index (low-priced index) but also relative to the old index (high-priced index). Turning to the difference in the change in betas, $\Delta \beta_2 - \Delta \beta_1$, we find that none of these differences are statistically or economically significant. In other words, the splitting stocks move more with both the old index and the new index to the same extent. Thus, there is no evidence of excess comovement following stock splits. All of the apparent effect is attributable to an increase in the fundamental beta of these stocks. Interestingly, the flawed bivariate regressions appear to continue to show an increase in comovement with the new index, a discussion of which we defer to later in the next section.
5. Comovement and Momentum

It would be a remarkable coincidence if selecting samples based on S&P500 index additions and stock splits was independently choosing stocks whose betas increase after the event. However, as it turns out, these two samples have something in common. The stocks in both samples have abnormally good performance before the event. This phenomenon is well known for stock splits—only companies whose stock price goes up split their stocks—but it is also intuitive for index additions—S&P is biased towards larger, better-performing stocks for inclusion in their flagship index, holding other criteria constant. The questions are (1) whether selecting on positive past performance can explain the beta increases that are consistent with the initial empirical results in Section 4, and (2) whether controlling for this effect eliminates the appearance of excess comovement. We look at the former question in Section 5.1 and the latter in Section 5.2.

5.1 Momentum and Beta

In examining changes in beta following periods of good performance, we follow the momentum methodology described in Section 3. While our focus is on winners, we report the winner and loser stock betas beginning 2 years before the holding period and continuing up to 2 years after the beginning of the holding period. The results, in Table 3 and Figure 1, show that betas of winner stocks increase dramatically during the formation period and continue to increase during the holding period. They stabilize thereafter for a few months and begin to decline. Specifically, we find that betas of winner stocks increase from 1.043 to 1.232 (a statistically significant change of 0.189) from Year-1 to Year 0, and from 1.023 in Year-2 to 1.232 in Year 0, a statistically and economically significant increase of 0.209. The betas continue to increase further during the holding period to 1.334 (a statistically significant change of 0.101) from Year 0 to Year+1 before declining to 1.218 in Year+2.

10 These tests require a long trading period potentially leading to a survivorship bias. The results, however, are virtually unaffected even when shorter periods are used.
This pattern of consistently increasing betas for stocks with high past returns has the potential to explain the results in Section 4. The betas of the stocks in the sample increase around the event in question, and therefore they comove more with all stocks after the event, both stocks in the group they are joining and stocks in the group they are leaving.

5.2 Comovement with Momentum Matched Firms (and Dimson’s betas)

For the analysis in this subsection, we make two adjustments in order to better assess the magnitude of excess comovement, if any, present in the data. First, because we are using daily data, nonsynchronous trading may be interfering with our ability to get accurate regression coefficients. To the extent that stocks do not all trade simultaneously at the end of each day, the observed return on a stock will be potentially correlated with leads and lags of the returns on a given portfolio. The correct adjustment for this effect in order to uncover the true regression coefficient is to sum the coefficients in a regression which includes these leads and lags (Dimson (1979)). Throughout the analysis in this section, we use two leads and lags for all portfolio returns used as independent variables.

Second, following up on the Section 5.1 results where we found evidence of increasing betas in momentum stocks, we also compare comovement of sample stocks with a matched sample that exhibits similar momentum characteristics. Barberis, Shleifer, and Wurgler (2005) use a sample of firms matched by size and industry, but do not control for momentum, which appears to be the critical factor due to the beta patterns associated with winner stocks. Consequently, for each addition, we select a matched firm from the same size decile that is not a member of the S&P 500 index and is closest in terms of lagged 252-day return to the added firm at the time of inclusion. Like Barberis, Shleifer, and Wurgler (2005), Green and Hwang (2009) construct a sample matched by size and industry without controlling for momentum. The matched sample that we use in this paper for stock splits controls for both size and momentum. For each stock split, we first select a group of firms from the high-priced portfolio that fall in the same size decile. Thereafter, we choose firms that are closest to the splitting

11 In results not presented here, requiring the matched firm to be from the same industry as the sample firm does not change the results.
firm by momentum. The matched firm is the one that comes closest in price and momentum to the sample firm within the same size decile.

Tables 4 and 5 present the results for the S&P500 index addition and stock split samples, respectively. In both cases, Panel A provides the univariate regression results, while those for the bivariate regression are reported in Panel B. Within each panel, we first present the results for the sample of event stocks. These results are comparable to those in Tables 1 and 2, except that we now use the Dimson adjustment to estimate the coefficients. We then provide the estimation results for the matched sample. Finally, we show the difference between the original and momentum-matched samples.

For the S&P500 index additions, the Dimson adjustment alone generally accounts for more than 50% of the effect that appears in the original analysis. For example, $\Delta \beta_2$ in the most significant sub-period (1988-2000) drops from 0.214 to 0.078. Looking at the differences between the coefficient changes across regressions, $\Delta \beta_2 - \Delta \beta_1$, only in this same sub-period is the coefficient statistically positive with a value of 0.129 and a t-statistic of 2.55. However, a similar result holds for the matched sample in 1988-2000. The strength of the momentum effect produces a value of 0.111 with a t-statistic of 2.23 for matched firms that are not added to the index. Consequently, in this sub-period and all the others, there is no single difference above 0.020 between the original and matched samples. To put it succinctly, there is absolutely no evidence of any excess comovement once we control for the momentum effect. For the bivariate regressions, which we consider unreliable, the same basic results of no excess comovement hold.

For stock splits, we have already established that even the original sample exhibits little or no evidence of excess comovement when comparing the univariate regression results across low-priced stocks (the new group) and high-priced stocks (the old group). Nevertheless, it is still worthwhile looking briefly at the results with Dimson betas for a momentum matched sample. Though we estimate Dimson betas for uniformity, we don’t anticipate Dimson betas making a significant difference because non-synchronous trading is unlikely to be different for the high-priced and low-priced groups. On the other hand, almost all splitting stocks are likely
to be momentum stocks so a properly matched sample should also exhibit similarly high changes in betas.

The basic results in Table 5 are affected little by the Dimson adjustment—comovement with both portfolios increases after the split by similar amounts. Not surprisingly, the same phenomenon shows up in the matched sample, although it is smaller than in the original sample. When taking differences, the values are economically very small and predominantly statistically insignificant. Similar results obtain for the bivariate regressions and excess comovement is not evident in any sub-period except during 1991-2004. We reason that the inconsistency in bivariate regression results across different sub-periods and the unreliability of results evident from the model in Section 2 make the univariate regression results more compelling. Thus, there is no meaningful evidence of excess comovement.

6. Robustness Checks

We reconfirm the baseline results on comovement in several ways. In particular, we examine whether there is comovement among the momentum stocks. Second, we repeat our analysis with weekly data and for index deletions.

6.1 Comovement among Momentum Stocks

Another approach to cement the connection between comovement and momentum is to study whether winner stocks exhibit comovement irrespective of whether or not there is an asset class shift from one group to another. Since the market betas of winner stocks increase and betas of loser stocks decrease marginally, winner stocks should comove with other winner stocks and loser stocks should comove with other loser stocks.

To conduct this test, we form a winner index based on stocks identified as winners and a loser index based on stocks identified as losers in the 6-month formation period, which we call the first group of winners and losers. We create a second group of winner stocks based on 3-month returns after the initial formation period but include only new winner stocks, which are
not in the first group of winners or losers. Similarly, we create a second group of loser stocks based on 3-month returns, which are not in the first group of losers or winners. The process is repeated on a monthly rolling basis. We measure the comovement of the second group of winner stocks with the first group of winner stocks and first group of loser stocks in the 3-month period following the second formation period by regressing the returns of the second group of winner stocks on the returns of the first group of winner stocks and returns of the first group of loser stocks.

\[
R_{w2} = \alpha + \beta_{w1}R_{w1} + \beta_{l1}R_{l1} + e
\]

\[
R_{l2} = \alpha + \beta_{w1}R_{w1} + \beta_{l1}R_{l1} + e
\]

We expect the coefficient to be larger on the first group of winner stocks than the first group of loser stocks. We repeat the process for the second group of loser stocks but with the same independent variables. In this case, we expect the magnitudes of the coefficients to be reversed: smaller on the first group of winner stocks and larger on the first group of loser stocks.

In results not reported here, we find that when the returns of the second group of winners is regressed on the first group of winners and the first group of losers, the coefficient on the first group of winners is significantly larger than the coefficient on the first group of losers. Similarly, when we regress the returns of the second group of losers on the same set of independent variables, the coefficient on the first group of losers is much larger than the coefficient on the first group of winners.

These results suggest that returns to recent winners are more likely to correlate with returns of past winners than past losers, and that returns to recent losers are more likely to correlate with returns of past losers than past winners. These results for momentum stocks are similar to the results reported by Barberis, Shleifer, and Wurgler (2005) and Green and Hwang (2009) and consistent with the notion that comovement may be another manifestation of beta changes due to momentum.

6.2 Weekly Data
Though Barberis, Shleifer, and Wurgler (2005) and Green and Hwang (2009) present evidence of comovement using daily, weekly and monthly data, their results are the strongest with daily data. Accordingly, the main results in the paper are based on daily data. Here we test the results with weekly data. Essentially, using weekly data has an effect similar to adding two leads and lags to the beta estimates, as we do above. Not surprisingly, the results are much weaker with weekly data than with daily data. Once a matched sample is used to control for changes in fundamental factors, there is no evidence of residual excess comovement in either univariate regressions or bivariate regressions.  

6.3 Index Deletions

Stocks are both added to and deleted from the S&P 500 index. Most index additions are at the discretion of the Index committee of Standard and Poor’s. However, index deletions may be voluntary or involuntary. The majority of index deletions are involuntary and occur because of structural changes such as a mergers, bankruptcies, or spinoffs. The remaining index deletions occur because a firm is no longer representative of the U.S. economy. In other words, the firm has become too small in size and is more appropriate for one of the smaller S&P indices or it belongs to an industry that S&P wants to down-weight due to its declining importance to the overall economy. Therefore, we expect evidence of comovement in index deletions to be much weaker, and we anticipate that a significant fraction of the comovement may be explained by non-synchronous trading.

We duplicate the analyses in Tables 1 and 4 for index deletions, and find results that are consistent with our results for index additions, that is, there is no excess comovement in stocks that are deleted from the S&P 500 index either in univariate or bivariate regression results.

7. Conclusion

---

12 Results are not reported in the interests of brevity.
Motivated by a simple model that captures the essence of the excess comovement hypothesis, we revisit the results of two well-known papers in the literature on comovement before and after S&P500 index additions (Barberis, Shleifer, and Wurgler, 2005) and stock splits (Green and Hwang, 2009). The model implies that looking at univariate regressions rather than bivariate regressions is more informative about the economic magnitude of the effect of interest, and, in particular, that the differences between the coefficients in univariate regressions on the returns of the group that the stock is leaving and the group that it is joining identify this effect. When we conduct this empirical exercise, the evidence points strongly to the conclusion that the existing results are due not to excess comovement but to changes in the comovement of stocks with fundamentals. These beta changes themselves are a feature common to winner stocks, an empirical phenomenon the documentation of which may be new to the literature. By making sure to measure these fundamental betas accurately, and controlling for this effect using a matched sample of winner stocks, we show that there is no longer any evidence of meaningful excess comovement from either an economic or statistical standpoint.
Appendix: Proofs

In the univariate regressions
\[ y_t = \alpha + \beta_1 x_{1t} + \epsilon_t \]
\[ y_t = \alpha + \beta_2 x_{2t} + \epsilon_t', \]
the probability limit of the slope coefficient estimates are
\[ \beta_1 = \frac{\text{cov}(y_t, x_{1t})}{\text{var}(x_{1t})}, \quad \beta_2 = \frac{\text{cov}(y_t, x_{2t})}{\text{var}(x_{2t})}. \]

Assume the driving processes for returns prior to the group switch are
\[ y_t = b_{y} f_t + c_{u} u_{1t} + e_{yt}, \]
\[ x_{1t} = b_{1} f_t + u_{1t} + e_{1t}, \]
\[ x_{2t} = b_{2} f_t + u_{2t} + e_{2t}, \]
\[ \text{var}(e_{u}) \equiv \sigma_{e1}^2, \quad \text{var}(u_{1t}) \equiv \sigma_{u1}^2, \quad \text{var}(f) \equiv \sigma_f^2, \]
and similarly after the group switch
\[ y_t = \tilde{b}_{y} f_t + \tilde{c}_{u} u_{2t} + e_{yt}, \]
\[ x_{1t} = \tilde{b}_{1} f_t + u_{1t} + e_{1t}, \]
\[ x_{2t} = \tilde{b}_{2} f_t + u_{2t} + e_{2t}, \]
\[ \text{var}(e_{u}) \equiv \sigma_{e2}^2, \quad \text{var}(u_{2t}) \equiv \sigma_{u2}^2, \quad \text{var}(f) \equiv \sigma_f^2, \]

Computing the coefficients prior to and after the switch of stock \( y \) from group 1 to group 2:
\[ \beta_1 = \frac{\text{cov}(b_{y} f_t + c_{u} u_{1t} + e_{yt}, b_{1} f_t + u_{1t} + e_{1t})}{\text{var}(b_{1} f_t + u_{1t} + e_{1t})} = \frac{b_{y} b_{1} \sigma_{f}^2 + c_{u} \sigma_{u1}^2}{\sigma_{e1}^2}, \]
\[ \beta_2 = \frac{\text{cov}(\tilde{b}_{y} f_t + \tilde{c}_{u} u_{2t} + e_{yt}, \tilde{b}_{1} f_t + u_{1t} + e_{1t})}{\text{var}(\tilde{b}_{1} f_t + u_{1t} + e_{1t})} = \frac{\tilde{b}_{y} \tilde{b}_{1} \sigma_{f}^2}{\sigma_{e2}^2}. \]
\[ \sigma_{e1}^2 = b_{2}^2 \sigma_{f}^2 + \sigma_{u1}^2 + \sigma_{e1}^2, \quad \sigma_{e2}^2 = \tilde{b}_{2}^2 \sigma_{f}^2 + \sigma_{u2}^2 + \sigma_{e2}^2. \]

Similarly,
\[ \beta_1 = \frac{b_{y} b_{2} \sigma_{f}^2}{\sigma_{x2}^2}, \quad \beta_2 = \frac{\tilde{b}_{y} \tilde{b}_{2} \sigma_{f}^2 + \tilde{c}_{u} \sigma_{u2}^2}{\sigma_{x2}^2}, \]
\[ \sigma_{x2}^2 = b_{2}^2 \sigma_{f}^2 + \sigma_{u2}^2 + \sigma_{e2}^2, \quad \sigma_{x2}^2 = \tilde{b}_{2}^2 \sigma_{f}^2 + \sigma_{u2}^2 + \sigma_{e2}^2. \]

Assuming the parameters other than stock \( y \)'s loadings on the fundamental factor and the non-fundamental group shocks are fixed across the 2 sub-periods, i.e.,
\[
\begin{align*}
\bar{b}_1 &= \bar{b}_1 \equiv b_1 & \bar{b}_2 &= \bar{b}_2 \equiv b_2 & \sigma_f^2 &= \sigma_f^2 \equiv \sigma_f^2 & \sigma_{ui}^2 &= \sigma_{ui}^2 \equiv \sigma_{ui}^2 & \sigma_{ei}^2 &= \sigma_{ei}^2 \equiv \sigma_{ei}^2
\end{align*}
\]

then
\[
\bar{\beta}_1 - \beta_1 = \frac{(\bar{b}_3 - b_3)\sigma_f^2 - c_1\sigma_{ui}^2}{\sigma_{x1}^2}
\]

\[
\bar{\beta}_2 - \beta_2 = \frac{(\bar{b}_3 - b_3)\sigma_f^2 + \bar{c}_2\sigma_{u2}^2}{\sigma_{x2}^2}
\]

If, in addition \(b_y = \bar{b}_y \equiv b_y\), then
\[
\bar{\beta}_1 - \beta_1 = -\frac{c_1\sigma_{ui}^2}{\sigma_{x1}^2} < 0
\]

\[
\bar{\beta}_2 - \beta_2 = \frac{\bar{c}_2\sigma_{u2}^2}{\sigma_{x2}^2} > 0
\]

Next, consider the bivariate regression:
\[
y_i = \alpha + \beta_{1b} x_{1i} + \beta_{2b} x_{2i} + \epsilon_i
\]

The probability limits of the coefficients are
\[
\beta = (X^T X)^{-1} (X^T Y) \quad \Rightarrow
\]

\[
\beta_{1b} = \frac{\text{cov}(y_i, x_{1i}) \text{var}(x_{2i}) - \text{cov}(y_i, x_{2i}) \text{cov}(x_{1i}, x_{2i})}{\text{var}(x_{1i}) \text{var}(x_{2i}) - \text{cov}(x_{1i}, x_{2i})^2}
\]

\[
\beta_{2b} = \frac{\text{cov}(y_i, x_{2i}) \text{var}(x_{1i}) - \text{cov}(y_i, x_{1i}) \text{cov}(x_{1i}, x_{2i})}{\text{var}(x_{1i}) \text{var}(x_{2i}) - \text{cov}(x_{1i}, x_{2i})^2}
\]

where the coefficients reflect a natural symmetry. It is convenient to rewrite these expressions in terms of the univariate coefficients defined above:

\[
\beta_{1b} = \frac{\text{cov}(y_i, x_{1i}) \text{var}(x_{2i}) - \text{cov}(y_i, x_{2i}) \text{cov}(x_{1i}, x_{2i})}{(1 - \text{corr}(x_{1i}, x_{2i})^2) \text{var}(x_{1i}) \text{var}(x_{2i})}
\]

\[
= \frac{1}{1 - \text{corr}(x_{1i}, x_{2i})^2} \left( \frac{\text{cov}(y_i, x_{1i})}{\text{var}(x_{1i})} - \text{corr}(x_{1i}, x_{2i}) \frac{\text{cov}(y_i, x_{2i})}{\text{var}(x_{1i}) \text{var}(x_{2i})} \right)
\]

\[
= \frac{1}{1 - \text{corr}(x_{1i}, x_{2i})^2} \left( \beta_1 - \text{corr}(x_{1i}, x_{2i}) \frac{\text{var}(x_{2i})}{\text{var}(x_{1i})} \beta_2 \right)
\]

\[
\beta_{2b} = \frac{1}{1 - \text{corr}(x_{1i}, x_{2i})^2} \left( \beta_2 - \text{corr}(x_{1i}, x_{2i}) \frac{\text{var}(x_{1i})}{\text{var}(x_{2i})} \beta_1 \right)
\]

As above, computing these values prior to and after the switch of stock \(y\) from group 1 to group 2:
\[
\beta_{ib} = \frac{1}{1 - \rho_{x_1,x_2}^2} \left[ \beta_1 - \rho_{x_1,x_2} \frac{\sigma_{x_2}}{\sigma_{x_1}} \beta_2 \right] \quad \beta_{2b} = \frac{1}{1 - \rho_{x_1,x_2}^2} \left[ \beta_2 - \rho_{x_1,x_2} \frac{\sigma_{x_1}}{\sigma_{x_2}} \beta_1 \right]
\]
\[
\bar{\beta}_{ib} = \frac{1}{1 - \rho_{x_1,x_2}^2} \left[ \bar{\beta}_1 - \rho_{x_1,x_2} \frac{\sigma_{x_2}}{\sigma_{x_1}} \bar{\beta}_2 \right] \quad \bar{\beta}_{2b} = \frac{1}{1 - \rho_{x_1,x_2}^2} \left[ \bar{\beta}_2 - \rho_{x_1,x_2} \frac{\sigma_{x_1}}{\sigma_{x_2}} \bar{\beta}_1 \right]
\]
\[
\rho_{x_1,x_2} = \frac{\text{cov}(\bar{x}_1, \bar{x}_2)}{\sqrt{\sigma_{x_1} \sigma_{x_2}}} \quad \text{cov}(\bar{x}_1, \bar{x}_2) = b_1 b_2 \sigma_f^2 + \text{cov}(\bar{e}_1, \bar{e}_2)
\]
\[
\rho_{x_1,x_2} = \frac{\text{cov}(\bar{x}_1, \bar{x}_2)}{\sqrt{\sigma_{x_1} \sigma_{x_2}}} \quad \text{cov}(\bar{x}_1, \bar{x}_2) = b_1 b_2 \sigma_f^2 + \text{cov}(\bar{e}_1, \bar{e}_2)
\]

Again assuming the parameters other than the weights on the non-fundamental group shocks are fixed across the 2 sub-periods,

\[
\beta_{ib} - \bar{\beta}_{ib} = \frac{1}{1 - \rho_{x_1,x_2}^2} \left[ (\beta_1 - \bar{\beta}_1) - \rho_{x_1,x_2} \frac{\sigma_{x_2}}{\sigma_{x_1}} (\beta_2 - \bar{\beta}_2) \right] > 0
\]
\[
\beta_{2b} - \bar{\beta}_{2b} = \frac{1}{1 - \rho_{x_1,x_2}^2} \left[ (\beta_2 - \bar{\beta}_2) - \rho_{x_1,x_2} \frac{\sigma_{x_1}}{\sigma_{x_2}} (\beta_1 - \bar{\beta}_1) \right] < 0
\]

If we further assume

\[
\sigma_{e_1}^2 = \sigma_{e_2}^2 = 0 \quad \bar{e}_1 = \bar{e}_2 = 1 \quad b_y = b_i = b_2 = 1
\]

then

\[
\beta_1 = \frac{\sigma_f^2 + \sigma_{u_1}^2}{\sigma_f^2 + \sigma_{u_1}^2} = 1 \quad \bar{\beta}_1 = \frac{\sigma_f^2}{\sigma_f^2 + \sigma_{u_1}^2}
\]
\[
\beta_2 = \frac{\sigma_f^2}{\sigma_f^2 + \sigma_{u_2}^2} \quad \bar{\beta}_2 = \frac{\sigma_f^2 + \sigma_{u_2}^2}{\sigma_f^2 + \sigma_{u_2}^2} = 1
\]
\[
\rho_{x_1,x_2} = \frac{\sigma_f^2}{\sqrt{\sigma_f^2 + \sigma_{u_1}^2} \sqrt{\sigma_f^2 + \sigma_{u_2}^2}}
\]
\[
\beta_{ib} = \frac{1}{1 - \rho_{x_1,x_2}^2} \left[ 1 - \rho_{x_1,x_2} \frac{\sigma_f^2}{\sigma_f^2 + \sigma_{u_1}^2} - \frac{\sigma_f^2}{\sigma_f^2 + \sigma_{u_2}^2} \right] = \frac{1}{1 - \rho_{x_1,x_2}^2} \left[ 1 - \rho_{x_1,x_2}^2 \right] = 1
\]
\[
\beta_{2b} = \frac{1}{1 - \rho_{x_1,x_2}^2} \left[ \frac{\sigma_f^2}{\sigma_f^2 + \sigma_{u_2}^2} - \frac{\sigma_f^2}{\sigma_f^2 + \sigma_{u_1}^2} \frac{\sigma_f^2}{\sqrt{\sigma_f^2 + \sigma_{u_1}^2} \sqrt{\sigma_f^2 + \sigma_{u_2}^2}} \right] = 0
\]

Similarly,

\[
\bar{\beta}_{ib} = 0 \quad \bar{\beta}_{2b} = 1
\]
References


Von Drathen, C., 2013. Is there really excess comovement? Causal evidence from FTSE1000 index turnover, working paper, LSE.
We estimate the univariate and bivariate regressions

\[ y_i = \alpha + \beta_1 x_{1i} + \varepsilon_i, \]
\[ y_i = \alpha + \beta_2 x_{2i} + \varepsilon_i, \]
\[ y_i = \alpha + \beta_{16} x_{1i} + \beta_{26} x_{2i} + \varepsilon_i, \]

for a sample of stocks that are added to the S&P 500 index from 1962 through 2011. The pre-event estimation period covers a one year window ending at the end of the month preceding announcement, while the post-event period covers the one year window starting the month after the effective date of index change. \( x_{1i} \) and \( x_{2i} \) are returns to non-S&P 500 index and S&P 500 index at time \( t \). Panel A reports the univariate regression results, and Panel B reports the bivariate regression results. In each cell, the first number is the mean and the second number is the corresponding t-statistic, where standard errors are clustered by month.

### Panel A. Univariate Regressions

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<th>( \bar{\beta}_1 )</th>
<th>( \Delta \beta_1 )</th>
<th>( \sigma_{\varepsilon_1}^2 \Delta \beta_1 ) (x 10^6)</th>
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<th>( \bar{\beta}_2 )</th>
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<th>( \sigma_{\varepsilon_2}^2 \Delta \beta_2 ) (x 10^6)</th>
<th>( \frac{\sigma_{\varepsilon_2}^2 \Delta \beta_2}{\sigma_{\varepsilon_1}^2 \Delta \beta_1} ) (x 10^6)</th>
<th>( \frac{\sigma_{\varepsilon_2}^2 \Delta \beta_2 - \sigma_{\varepsilon_1}^2 \Delta \beta_1}{\sigma_{\varepsilon_2}^2 \Delta \beta_1} ) (%)</th>
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Panel B. Bivariate Regressions

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<th>$\beta_{2b}$</th>
<th>$\bar{\beta}_{2b}$</th>
<th>$\Delta \beta_{2b}$</th>
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<td>0.339</td>
<td>0.648</td>
<td>0.235</td>
<td>0.278</td>
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Table 2: Stock Splits

We estimate the univariate and bivariate regressions

\[ y_i = \alpha + \beta_1 x_{1i} + \varepsilon_i; y_i = \alpha + \beta_2 x_{2i} + \varepsilon_i \]

for a sample of 2-for-1 stock splits from 1962 through 2011. Our sample include all ordinary common stock two-for-one splits with a pre-split price of $10 or greater during our sample period. \( x_{1i} \) and \( x_{2i} \) are return to a portfolio of high priced stocks whose price belongs to \([3p/4, 5p/4]\) and low price stocks with prices within \([1p/4, 3p/4]\) at time \( t \), where \( p \) is the pre-split price before effective date of split. The pre-event (post-event) window is defined as the one year ending (beginning) one month before (after) the split date. Panel A reports the univariate regression results and Panel B reports the bivariate regression results. In each cell, the first number is the mean and the second number is the corresponding t-statistic, where standard errors are clustered by month.

| Sample Period | nobs | \( \beta_1 \) | \( \overline{\beta}_1 \) | \( \Delta \beta_1 \) \((\times 10^6)\) | \( \sigma^2 \Delta \beta_1 \) | \( \beta_2 \) | \( \overline{\beta}_2 \) | \( \Delta \beta_2 \) \((\times 10^6)\) | \( \sigma^2 \Delta \beta_2 \) | \( \sigma^2 \Delta \beta_2 - \sigma^2 \Delta \beta_1 \) \((\times 10^6)\) | \( \Delta \beta_1 \) \((\%)\) | \( \Delta \beta_1 \) \((\times 10^6)\) |
|---------------|------|----------------|-----------------|----------------|----------------|----------------|----------------|----------------|----------------|----------------|----------------|----------------|----------------|
| 1962-1970     | 543  | 1.044          | 1.146           | 0.101          | 3.758          | 1.178          | 1.251          | 0.073          | 2.842          | 3.758          | -0.916         | -0.157         | -0.028         |
| 1971-1990     | 2,349| 0.736          | 0.929           | 0.193          | 16.230         | 0.847          | 1.043          | 0.196          | 13.884         | 16.230         | -2.347         | -0.441         | 0.002          |
| 1991-2004     | 2,478| 0.798          | 1.014           | 0.216          | 31.892         | 0.937          | 1.186          | 0.248          | 29.443         | 31.892         | -2.449         | 0.162          | 0.033          |
| 2005-2011     | 408  | 1.088          | 1.236           | 0.148          | 16.429         | 1.073          | 1.202          | 0.129          | 12.888         | 16.429         | -3.541         | -0.430         | -0.019         |
|               |      | 34.725         | 29.560          | 4.558          | 4.243          | 36.786         | 34.122         | 3.856          | 3.561          | 4.243          | -1.929         | -1.232         | -0.966         |
| 1971-2011     | 5,235| 0.793          | 0.993           | 0.200          | 23.659         | 0.907          | 1.123          | 0.215          | 21.171         | 23.659         | -2.488         | -0.154         | 0.015          |
| 1962-2011     | 5,778| 0.816          | 1.007           | 0.191          | 21.789         | 0.933          | 1.135          | 0.202          | 19.448         | 21.789         | -2.340         | -0.155         | 0.011          |
|               |      | 67.681         | 74.289          | 18.599         | 9.559          | 70.137         | 71.408         | 18.754         | 10.748         | 9.559          | -2.742         | -1.900         | 2.137          |
Panel B. Bivariate Regressions

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<td>0.989</td>
<td>0.073</td>
<td>0.075</td>
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Table 3: Beta Changes and Momentum

At the end of each June from 1962 through 2010, stocks with a price of at least $10 that do not fall into the bottom size decile of NYSE stocks are assigned into 10 momentum deciles based on their cumulative returns over the preceding 252 days. We estimate betas for each stock based on a rolling window of 252 days from two years before formation of momentum portfolios through two years after formation, and compare beta changes for both the top and bottom two momentum portfolios. Thus, betas for years -2 and -1 are estimated over rolling windows ending 504 and 252 trading days before portfolio formation, respectively. Post-momentum portfolio formation years allow for a 21-trading day skip, and are estimated over 252 days ending 273 and 525 trading days after portfolio formation. In each cell, the first number is the time series average of the mean, and the second number is the corresponding t-statistic.

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<th>Year -1</th>
<th>Year 0</th>
<th>Year 1</th>
<th>Year 2</th>
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Table 4: S&P Additions with Matched Sample and Dimson Adjustments

We estimate the univariate and bivariate Dimson (1979) regressions for a sample of stocks that are added to the S&P 500 index from 1962 through 2011 and for a portfolio of matched firms. The pre-event estimation period covers a one year window ending at the end of the month preceding announcement, while the post-event period covers the one year window starting the month after the effective date of index change. \(x_{1t}\) is return to non-S&P 500 index at time \(t\), while \(x_{2t}\) is return to the S&P 500 index at time \(t\). The match firm for each addition is identified as the one with closest momentum from the same size decile as the addition firms. The Dimson beta is defined as a simple sum of the lag, concurrent, and lead coefficients from the following regressions with two leads and lags. In each cell, the first number is the mean and the second number is the corresponding t-statistic, where standard errors are clustered by month.

\[
y_t = \alpha + \sum_{s=2}^{2} \beta_1^s x_{1,t+s} + \epsilon_t
\]

\[
y_t = \alpha + \sum_{s=2}^{2} \beta_2^s x_{2,t+s} + \epsilon_t
\]

\[
y_t = \alpha + \sum_{s=2}^{2} \beta_1^s x_{1,t+s} + \sum_{s=2}^{2} \beta_2^s x_{2,t+s} + \epsilon_t
\]
## Panel A. Univariate Regressions

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Table 5: Stock Splits with Matched Sample and Dimson Adjustments

We estimate the univariate and bivariate Dimson (1979) regressions for a sample of 2-for-1 stock splits from 1962 through 2011. Our sample include all ordinary common stock two-for-one splits with a pre-split price of $10 or greater during our sample period. $x_1$, and $x_2$, are return to a portfolio of high priced stocks whose price belongs to $[3p/4, 5p/4]$ and low price stocks with prices within $[1p/4, 3p/4]$ at time $t$, where $p$ is the pre-split price before effective date of split. The pre-event (post-event) window is defined as the one year ending (beginning) one month before (after) the split date. The Dimson beta is defines as a simple sum of the lag, concurrent, and lead coefficients from the following regressions with two leads and lags. In each cell, the first number is the mean and the second number is the corresponding $t$-statistic, where standard errors are clustered by month. In each cell, the first number is the mean and the second number is the corresponding $t$-statistic, where standard errors are clustered by month.

\[
y_t = \alpha + \sum_{s=-2}^{2} \beta_1^s x_{1,t+s} + \epsilon_t \\
y_t = \alpha + \sum_{s=-2}^{2} \beta_2^s x_{2,t+s} + \epsilon_t \\
y_t = \alpha + \sum_{s=-2}^{2} \beta_1^s x_{1,t+s} + \sum_{s=-2}^{2} \beta_2^s x_{2,t+s} + \epsilon_t
\]
## Panel A. Univariate Regressions

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## Panel B. Bivariate Regressions

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We estimate market betas of winner and loser stocks defined as the top and bottom deciles of stocks sorted on past 12-month returns, skipping the most recent month, as in Jegadeesh and Titman (2001). These betas are estimated over rolling windows of 252 days (1 year).